Was There A British House Price Bubble?
Evidence from a Regional Panel*

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Abstract: This paper investigates the bubbles hypothesis with a dynamic panel data model of British regional house prices between 1972 and 2003. The model consists of a system of inverted housing demand equations, incorporating spatial interactions and lags and relevant spatial parameter heterogeneity. The results are data consistent, with plausible long-run solutions and include a full range of explanatory variables. Novel features of the model include transaction cost effects influencing the speed of adjustment, and interaction effects between an index of credit availability and real and nominal interest rates. No evidence for a recent bubble is found.
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1. Introduction

This paper addresses the question of whether there was recently a bubble in UK house prices, which we take to mean a systematic but temporary deviation of house prices from fundamentals. For example, Barrell (2004) and the OECD (2005) suggest that UK house prices were overvalued by 30% or more in 2003-4, while Nickell (2005) rejects the bubble hypothesis.[1] Much of the debate about house price bubbles focuses on charts of ratios of house prices to income or rents, or mortgage payment to income ratios. These ratios are not very informative about the presence of absence of bubbles, because they ignore a range of other important factors, including demographic and population changes, house-building, credit conditions, and other asset prices. Systematic and comprehensive models are needed to distinguish fact from fiction in assessing the bubble hypothesis.

This paper presents a dynamic equilibrium-correction equation system for annual house prices in nine regions of Britain to exploit the richness of regional data. The model, estimated over the period 1972 to 2003, consists of a system of inverted housing demand equations with the predetermined regional housing stock appearing as an explanatory variable. The demand shifters include regional incomes, real and nominal interest rates, and demographics. The model incorporates spatial interactions and lags as well as relevant spatial parameter heterogeneity.

A major advantage of the inverted demand function approach is that we have strong priors regarding the values of the key long run elasticities, corresponding to the "central estimates" set out in Meen (1996, 2002) and Meen and Andrews (1998) inter alia. We estimate our model for 1972-1996 and forecast for 1997-2003 finding no evidence of systematic under-prediction for 1997-2003, or for 2000-2003. Alternative forecast scenarios for 2004-2010 also suggest that only very negative shifts in fundamentals, that few currently expect, are likely to lead to significant falls in nominal prices.

[1] The OECD appears to base their conclusions on two pieces of evidence. The first is a comparison of house price to rent ratios and a backward looking measure of the 'user cost' of housing, which combines interest costs and expected house price appreciation. The second piece of evidence used is the IMF (2005) house price equation. Nickell (2005), a member of the Bank of England's Monetary Policy Committee, argues that a shift in fundamentals occurred.

Our results have relevance outside the UK where there has also been a great deal of discussion of the possibility that there have been house price bubbles (see, inter alia, Case and Shiller, 2003, Himmelberg, Mayer and Sinai, 2005, and OECD, 2005). Credit market liberalization has occurred in other countries, as well as the UK. Our quantification of its effects, including shifting the relative role of real and nominal interest rates, is thus of wider interest. While, as discussed below, there are a number of US house price studies at the Metropolitan Area level, spatial interactions of the type we consider have usually been neglected. Further, in the US, borrowers increasingly choose adjustable-rate mortgages (Campbell and Cocco, 2003), and planning constraints have become more pervasive (Glaeser, Gyourko and Saks, 2005). In these respects, US housing markets now more resemble those of the UK than was once the case.

It is beyond the scope of this paper to survey the large theoretical literature on asset price bubbles, but we take a few examples. Bubbles may occur as a result of agency considerations in the banking sector (Allen and Gale, 2000) or differences between public and private information in games of incomplete information (Morris and Shin, 2002). Moreover, limits to arbitrage due to short-selling constraints may prevent rational traders from driving noise traders out of the market (De Long, Schleifer, Summers and Waldmann, 1990) and this, as well as the lack of deep pockets, is likely to make for inefficiencies in the housing market (Stein, 1995). It is interesting to note that policymakers and markets are now paying attention to the possibility of alleviating such constraints (Shiller, 2004, and Goetzmann et al. 2003).

For the US housing market, there is a wide range of opinion on whether or where there are bubbles. Important contributions to this debate have been made by Brunnermeier and Juilliard (2006) and Himmelberg et al. (2005) inter alia. Brunnermeier and Juilliard examine the effect of money illusion on house prices. Following Modigliani and Cohn (1979), they suggest that falling inflation and nominal interest rates (holding real interest rates fixed) leads to mispricing of housing since people wrongly attribute a decrease in inflation to a decline in the real interest rate and consequently underestimate the real cost of future mortgage payments. They argue that this mispricing explains a large and significant proportion of the run-up in UK and US real house prices. They argue that the effect of the nominal interest rate on house prices cannot be explained by the so-called ‘tilt’ effect, which arises because inflation shifts the real burden of mortgage payments towards the earlier years of the financing contract.

Himmelberg et al. (2005) argue that conventional metrics, such as house price to rent ratios, are misleading because they fail to account for long-run trends in real interest rates that have made housing rather more affordable in recent years. Their analysis of 46 Metropolitan Areas between 1980 and 2004 concludes that, although the late 1980s bubble shows up clearly in the data, the current high level of US house prices does not look overvalued once low long-term real interest rates are taken into account. Duca (2005) reaches a slightly more ambiguous position in an analysis of the relationship between home price to rent ratios and real mortgage rates; suggesting that prices were 11½ percent above their equilibrium values in the second quarter of 2005 - not enough to qualify as a bubble. At a regional level, however, Duca makes the point that price to rent ratios are further above their estimated equilibrium values in coastal cities with tough zoning regulations.

Regional house price models have many advantages provided good data exist. Regional data vary idiosyncratically and are thus more informative about the determinants of house prices than national data (see Figures 1 and 2), yielding more accurately estimated parameters. A regional house price model is also necessary to address regional issues such as the "ripple effect", whereby house prices in Greater London tend to lead prices in the South East and, with longer lags, the rest of Britain. Regional housing affordability, especially in Greater London and the South East, has become a major policy issue, raising the question of what the medium term impact will be of increasing the number of new homes to be built in different regions.

The outline of the paper is as follows. In Section 2, we outline the standard or textbook model of house prices which is the starting point of our research and extend it to the regional dimension. We also briefly review the recent literature on modelling regional house prices in the UK. We set out and present the results of estimating our regional house price model in Section 3. We illustrate and discuss our results further in Section 4. In Section 5 we use our results to address the issue of whether there has been a bubble in the British housing market. Our conclusions are set out in Section 6. The data used in the paper are described in the Appendix.

2. Modelling Aggregate and Regional House Prices

This Section outlines the standard or textbook model of national house prices and then lays out the basic characteristics of a system of regional equations. This is followed by a brief review of research on regional house prices in the UK.

(a) The standard model

The standard model of the housing market consists of three equations - a demand equation which (given the housing stock, real incomes, interest rates and so forth) largely determines house prices in the short run; a supply equation which determines the supply of new houses and an equation showing how the stock of houses changes over time as new houses are completed. The house price equation is derived from the demand for housing services by inverting and rearranging the demand equation, so that the dependent variable is house prices as opposed to the quantity of housing services or the housing stock. This is the most common form of house price equation in the international literature. For example, see Buckley and Ermisch (1982), Mankiw and Weil (1989), Meen (1990, 1996 and 2000), Muellbauer and Murphy (1994, 1997), Muth (1989) and Poterba (1984, 1991).

A simplified version of our house price equation can be derived as follows. A log-linear equation for the demand for housing services, which is assumed proportional to the housing stock, may be specified as:

(1) $\ln (hs/pop) = \alpha \ln (y/pop) - \beta \ln r_h + \ln d$

where hs is the housing stock, pop is population, y is real income, r_h is the real rental cost of housing and d represents other factors, such as demography, which shift the demand for housing. The α and β coefficients are the income and price elasticities of the demand for housing services. The international literature suggests that the income elasticity α is between ½ and 1 in cross section data and as high as 1¼ in time series data, and, that the price elasticity β is about ½, (see Meen (1996) and Meen and Andrews (1998) for example).

In the UK, the real rental is not directly observed since the private rental sector is small and not representative of the overall housing stock. However, in equilibrium, the rental equals the real user cost of housing. In the simplest case, the user cost is just:

$$(2) \qquad hp.(r_a + m + t_h - hp^e / hp) \equiv hp.uc_h$$

where hp = real price of houses, r_a = tax adjusted real interest rate, m = rate of expenditure on maintenance and repair etc., t_h = net rate of tax on housing and hp^e / hp = expected rate of appreciation of real house prices. Thus uc_h is the user cost of housing, expressed as a proportion of the price of the house. In practice, the main drivers of the user cost are the mortgage rate and the expected rate of inflation of house prices.

The inverted demand curve, obtained by substituting the user cost (2) for the rental in (1) and inverting, is then:

$$(3) \qquad \ln hp = \alpha/\beta \; \ln (y/pop) - 1/\beta \; \ln (hs/pop) - \ln uc_h + 1/\beta \; \ln d$$

House prices are positively related to real per capita incomes y/pop, negatively related to the per capita housing stock hs/pop and the user cost of housing uc_h and positively related to other variables that increase the demand for housing. Using central estimates of 1 for α, the income elasticity of housing demand, (3) may be re-written in terms of real income per house, the user cost and other demand shifters:

$$(3') \qquad \ln hp = \theta \ln (y/hs) - \ln uc_h + 1/\beta \; \ln d$$

where $\theta = \alpha/\beta$. This may be estimated as:[2]

$$(4) \qquad \ln hp = \beta_0 + \theta (\ln y - \ln hs) - \beta_1(r^a + m + t_h - \Delta \ln hp^e) + \beta_2 \ln d + u$$

Equations similar to (4) may also be derived from an explicit multi-period utility maximization problem. Income y is then a measure of permanent income or some combination of physical and financial wealth and current and future real income. At the core of our regional house price equations, is a long run equation very similar to equation (4).

[2] In practice, the level of uc_h is used in place of ln uc_h, since the user cost of housing is often negative, at least in the UK.

In practice, estimated versions of (4) are invariably dynamic - they include lagged house prices and lagged explanatory variables on the right hand side of (4) and often include an equilibrium correction term. Many of the modelling choices, such as the choice of proxies or selection of lag lengths, are largely data determined.

The unobserved $\Delta \ln hp^e$ term in the user cost in (4) has to be proxied in some fashion. For example, the expected capital gains may be proxied by lagged capital gains or the fitted value from a simple regression on predetermined variables. Sometimes, the interest rate and capital gains components of the user cost appear separately in the equation, with a larger coefficient on the interest rate term, Meen (1996, 2002). Kearl (1979) argued for the inclusion of nominal interest rates because of 'tilt' or 'front-end-loading', discussed further below. In UK studies, proxies for credit conditions or mortgage rationing are generally significant as well. Regional house price models, which we now turn to, are yet more complicated.

(b) The basics of regional house price models

Regional house price models are not just national house price models with regional data substituted for national data. A regional model also includes spatial lags (spillover or contiguity effects), spatial errors and possibly spatial coefficient heterogeneity. Consider a two region economy (r = 1,2) with log-linear, housing demands given by a simplified version of (1) with a unit income elasticity, and incorporating also the relative house price in the two regions:

(5) $$\ln hs_r = -\beta_r \ln hp_r - \gamma_r (\ln hp_r - \ln hp_s) + \ln y_r + z_r$$

for regions r,s = 1,2 and where z_r represents other demand shifters.[3] The own price elasticity of demand is $\beta_r + \gamma_r$ holding the price in the other region fixed, and the cross price elasticity is γ_r. Solving the two equations for $\ln hp_r$ yields

(6) $$\ln hp_r = [\ (\ln y_r - \ln hs_r + z_r) + \gamma_1/(\beta_2 + \gamma_2).(\ln y_s - \ln hs_s + z_s)\ ] / \kappa$$

where $\kappa = [\beta_1 + \gamma_1 - \gamma_1\gamma_2/(\beta_2 + \gamma_2)]$. This can also be parameterised in terms of region r and national variables.[4] Adding i.i.d. random error terms to (5), generates spatial

[3] Meen (2001) discusses the relationship between national, regional and other spatial house prices models.

errors (contemporaneous correlation in the random error terms) in (6) and any equation derived from (6). One can then add lagged adjustment, while the demand shifters or z_r's include real and nominal interest rate effects, lagged appreciation in both regions and other factors discussed further below.

Comprehensive reviews of the regional house price literature for the UK can be found in Muellbauer and Murphy (1994) and Meen and Andrews (1998), so we need only summarize the small number of papers which have not been reviewed elsewhere. Meen and Andrews (1998) and Meen (2001) suggest that any valid model of regional house prices should possess the following features: (i) it should be data consistent; (ii) It should incorporate spatial lags and errors; (iii) it should capture spatial patterns of coefficient heterogeneity; (iv) the implied estimates of the income and price elasticities of housing demand should be plausible; (v) it should include the full range of variables found to be important in national house price models; (vi) the implications for housing efficiency should be clear; (vii) it should be capable of explaining the ripple effect, and (viii) the relative importance of demographic, as opposed to economic determinants, should be clear. Unfortunately, the majority of the UK regional house price models, surveyed in Muellbauer and Murphy (1994), Meen and Andrew (1998), do not satisfy many of these criteria. The recent empirical literature is no better and tends to focus on statistical (i.e. unit root and cointegration) issues, either (i) include no explanatory variables apart from lagged or leading region house prices or (ii) only consider a small set of explanatory variables and/or (iii) use non-structural models which are difficult to interpret. As a result, most of these papers say little about the economic determinants of regional house prices or the likelihood of bubbles.

Drake (1995) uses a time varying parameter model to look for evidence of regional convergence in house prices and the ripple effect. Cook (2003, 2005) uses various unit root and stationarity tests to check for a stable ripple effect from London house prices. Wood (2003) also focuses on the ripple effect and tries to identify whether regional house prices have moved in a way consistent with the ripple effect. His results are mixed. However, his model of national house prices is extremely naïve –

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[4] For example, if the β's and γ's are all equal to ½ (a reasonably plausible value), then $\ln hp_r = {}^4/_3\, [(\ln y_r - \ln hs_r + z_r) + ½.(\ln y_s - \ln hs_s + z_s)]$ is approximately proportional to $½.(\ln y_r - \ln hs_r + z_s) + (\ln y - \ln hs + z)$ where y is national income and regional shares in national income are approximately constant etc. In this equation, the weight on own region income y_1 is half the weight on national income y.

national house prices are only determined by average earnings and the real interest rate in the long run. Meen (1999) suggests that income differences and spatial lags do not fully explain the ripple effect. Differences in structures between regions, which show up as coefficient heterogeneity in a model, are also important, he argues. The empirical results back up these claims.

We are not aware of a US literature on *regional* house price determination, though there are a number of studies of house prices for Metropolitan Areas (MAs). For example, Follain and Velz (1995) and Hwang and Quigley (2005) respectively study price movements in panels of 22 and 74 MAs. Abraham and Hendershott (1996), Malpezzi (1999) and Capozza et al. (2004) also examine the equilibrium correcting behaviour of house prices in MAs. However, spatial issues, which might, for example, arise through location choices by households resulting in relative price effects captured by the parameter γ_r in equation (5), are not addressed in any of these studies.

3. An Outline of the Regional House Price Model

We model real house prices in nine regions of Great Britain between 1972 and 2003, estimating both for 1972-1996 and 1972-2003 to check for parameter stability.[5] Estimating to 1996 and forecasting 1997-2003 is also an important part of the check for bubbles. Figure 1 shows log real regional house prices for Greater London, the West Midlands and the North. It suggests that the regions experienced broadly comparable long run movements. Greater London is considerably more expensive than the other regions.[6] Figure 2 shows the same information in terms of log changes, which allows the heterogeneity in movements to be seen more clearly. The leading role of Greater London and the tendency of house prices in the North to lag further behind those in the West Midlands are clear.

[5] The data are mix-adjusted house price indices scaled by the national consumer expenditure deflator. The eight regions are the North (NT), Yorkshire and Humberside (YH), East Midlands (EM), West Midlands (WM), Greater London (GL), the South (ST), the South West (SW), Wales (WW) and Scotland (SC). The choice of regions is determined by the need for consistent regional boundaries since the government switched from Standard Statistical Regions (SSR's) to Government Office Regions (GOR's) in the mid 1990's. The North region is the sum of the current North East and North West GOR's, which is the sum of the old North and North West SSR's. The South region is the sum of the South East and Eastern GOR's, which is the sum of the old Rest of South East (i.e. excluding Greater London) and East Anglia SSR's.

[6] Note that 1990 average second hand house prices in each region are used to scale the regional mix-adjusted indices.

--- Figures 1 and 2 About Here ---

We estimate a system of inverted housing demand equations using a two-step systems estimator, which takes account of contemporaneous spatial correlations. Step 1 uses SUR (seemingly unrelated regressions). Step 2 takes the covariance matrix estimated in step 1 as fixed in GLS.[7] The equations are non-linear with many cross-equation restrictions because of common parameters and interaction terms. However, some spatial coefficient heterogeneity is allowed for.

--- Tables 1 and 2 About Here ---

Our basic regional house price / inverted demand equation is set out in Table 1. We have annotated the equation to help the reader. The Greater London and South equations are similar to the equation set out above except that (i) some of the short run coefficients and time dummies are allowed to take different values and (ii) stock market effects are included.

The parameter estimates with robust standard errors are set out in Table 2. We imposed two restrictions, partly to save degrees of freedom, and partly to tie down the long run of the model when checking for non-linear interaction effects. The long run elasticity of house prices with respect to income and the housing stock θ is set to 1.6. This is close to the freely estimated value, and is near our prior and the central estimates in the literature. With an income elasticity of demand for houses of unity, also a statistically acceptable restriction, this implies an own price elasticity β of -0.625 when all regional prices vary in proportion.[8]

We model the log change in real house prices[9] $\Delta lrhp_r$ (the r subscript stands for region r) in each region (deflating by the consumer expenditure deflator) using lagged real log house price changes, real log income per house, and other variables. We check for inflation effects reflecting nominal inertia, and find that the rate of

[7] The 2 step procedure is more robust in small samples than maximum likelihood procedures, which maximize over the elements of the variance-covariance matrix.

[8] Of course, the price elasticity for region r, given prices in other regions, will be rather higher.

[9] Muellbauer and Murphy (1994) and some other researchers model relative house prices. Given a model for UK house prices, it does not matter greatly whether one models real or relative regional house prices. However, a model for real regional house prices is a good deal easier to interpret and can also be helpful in finding a coherent specification for a national house price model.

acceleration of prices, $\Delta^2 lpc$, conveniently summarises current and lagged inflation effects. We also checked for inflation volatility (see the Data Appendix for the definition), which we would expect to have a negative effect, since it will be associated with greater real and nominal interest rate uncertainty. However, it is closely negatively correlated with the credit conditions index discussed below, so that we could not find sufficiently stable coefficients.

The long-run solution is for $lrhp_r$, the real log level of house prices in region r. The key element in the long-run solution is the log of real personal disposable non-property income per house[10]. For region r, we write this as $lrynhs_r$ defined as log(real personal disposable non-property income in region r) - log(lagged housing stock in region r) - 0.7*log(lagged rate of owner-occupation in region r). The owner-occupation term suggests a modest spill-over from non-owner occupied supply onto the owner-occupied housing market (Muellbauer and Murphy, 1997). This means that one rented house added to the stock has around 30% of the effect on prices compared to the effect of adding a house for owner-occupation. The spill-over could well be larger, for example a weight of 0.5 on the log owner-occupation rate results in effectively the same fit but lower estimates of the region specific trends.

Population or the number of households is implicit in this formulation since income and the housing stock can both be put on a per capita basis, but the number of people or households just cancels out. However, we find that all regions are influenced not just by the own region value of income per house ($lrynhs_r$) but also by the value for Great Britain ($lrynhs_{GB}$), for reasons explained in Section 2(b). We can accept the restriction that the weights ($1 - w_0$ and w_0) on $lrynhs_r$ and $lrynhs_{GB}$ are equal to 0.33 and 0.67 respectively, and furthermore, there is surprisingly little evidence for heterogeneity across regions.[11]

[10] Data definitions and sources are set out in the Appendix. Note that we do not use regional personal income data from the Regional Accounts since it is biased (Cameron and Muellbauer, 2000). Instead we use regional earnings and employment data to modify the National Accounts personal income data to estimate personal non-property income on a regional basis.

[11] A value of 0.67 may initially appear large but some calculations suggest that this is not implausible. For example, given equation (6) consider the symmetric case with $\beta = 0.625$ and $\gamma = 0.5$. This implies a value for w_0 of around 0.67. See footnote 4 for a similar calculation. Interestingly, Hwang and Quigley (2005) for US metropolitan areas find an own price effect in the region of -1 to -1.2. But since there are no relative price terms in their equations, their own-price effect should reflect $\beta + \gamma$, which in our case is also estimated to be of the order of -1.

We present results both for a specification where the speed of adjustment (λ) is constant and for one in which it varies with the average Stamp Duty rate (sdr), an important ingredient of transactions costs.[12] Stamp Duty rates on houses have risen substantially since 1995. To our knowledge, this is the first time that the speed of adjustment of house prices has been conditioned on time varying transactions costs.

Each equation contains region specific intercepts and time trends. In all regions, these trends have positive coefficients. This may reflect trends in housing quality which do not show up in the housing stock data, or other trending variables such as income inequality, housing tenure (as noted above) and perhaps household size. We checked for income distribution effects, since space is a luxury good. However, income distribution changes are trend like and so hard to detect here.

Other key levels effects in the long-run solution include an index of credit conditions (cci)[13] and the interaction of this index with both the log nominal tax adjusted mortgage rate (labmr) and the real mortgage rate (rabmr). The inclusion of the real interest rate reflects its role in user cost. The nominal interest rate is included because, as Kearl (1979) argued, when nominal interest rates rise with inflation, leaving real rates unchanged, the real interest burden under standard mortgage contracts is more heavily tilted to or loaded on the first few years. However, as the liberalisation of credit markets has made refinancing easier, households have been better able to get round these near-term cash-flow constraints. We therefore interact the log nominal mortgage rate with cci, expecting to find a positive coefficient, weakening the negative effect of the nominal rate on house prices as cci rises. This interaction effect has a robust t-ratio of at least 3.3 in the results reported in Table 2. We use the log of the nominal rate because we regard the log of the debt service ratio as the relevant factor in a model for the log of house prices. A test of the log versus linear specification supports the former.

Brunnermeier and Julliard (2006) argue that inflation illusion rather than ‘tilt’ explains the significance of the nominal interest rate. As we do, they argue that the tilt hypothesis implies a temporal decline in the absolute size of the nominal interest

[12] In the UK, Stamp Duty is the tax you pay when you buy property or shares.

[13] The cci index was estimated by Fernandez-Corugedo and Muellbauer (2005) from data on 10 consumer credit and mortgage market indicators. It is intended to measure the shift in the credit supply function to UK households, especially since 1980.

rate effect, which they try to measure from moving window estimates. However, if anything, these show an increase in the effect, the opposite of our finding. The advantage of our approach, apart from the extra information content from regional data, is the use of a specific index of the shift in credit supply, cci. With such an index, we can control both for the direct effects of the credit supply shift on house prices, and for the interaction effect with the real mortgage rate. These controls are missing in Brunnermeier and Julliard (2006), which we regard as a mis-specification of their house price model. However, our findings do not preclude inflation illusion as a partial explanation of the nominal rate effect, merely that tilt is also important to it.

The same reasoning that suggests that nominal rate effects weaken as cci rises, suggest that real rate effects strengthen as cci rises, strongly confirmed by our empirical result. This result is consistent with findings for mortgage demand in Fernandez-Corugedo and Muellbauer (2005). The log nominal interest rate effect means that a reduction of rates from 5% to 4% has a stronger effect on house prices than a reduction from 10% to 9%. As well as level effects of the log nominal mortgage rate, we test for change effects. The short run nominal interest rate effects are also a little stronger in London and the South East. When modelling British house prices for periods including the 1970's, the inclusion of some measure of credit availability is necessary in order to obtain a correctly signed real interest rate effect.

Another level effect for which we checked is the log price of house prices in London relative to GB ($rlhp_{GL}$) whose coefficient we allow to vary by region. This proved insignificant.

We also investigated the role of the national mortgage default or possessions rate, as an indicator of fear or downside risk in the housing market, comparing it with an alternative, the average value over the previous four years of the negative return in the region's housing market, $rrhneg_r$.[14] The latter variable is significant and has the expected positive coefficient, eliminating the possessions rate variable from the

[14] The rate of return rrh_r is defined as the lagged nominal rate of house price appreciation in the region minus the tax adjusted mortgage interest rate/100 plus 0.03 to reflect the net benefit from owning a home. We define $rrhneg_r$ to equal rrh_r if rrh_r is negative, and zero otherwise.

model. This means that a history of negative returns depresses house prices for some time to come.[15] On its own, the national possessions rate would have been significant.

In the dynamics, the persistence of the previous year's house price growth rate is measured through a coefficient common to all regions. However, the relative weight attached to the own region ($\Delta lrhp_{r,-1}$), to regions contiguous to region r ($\Delta clrhp_{r,-1}$), and to Greater London ($\Delta lrhp_{GL,-1}$), varies by region. Our prior is that these weights lie in the interval 0 to 1. This hypothesis is acceptable when tested. The boundary restrictions on these weights are binding in all except the two Midlands regions and the North. Regions closer to London have a weight of 1 on London house price growth, reflecting the 'ripple effect' emanating from London. For the North and the two Midlands regions, there is mix of own region, contiguous region and/or London lagged price growth effects. For Wales and Yorkshire and Humberside, the contiguous region weight is 1 and for Scotland, the own region effect has a weight of 1. This is consistent with a longer lag from London in the more distant regions, with Scotland affected only by its own lagged house price change. However, it must be remembered that the derivation of equation (6) assumed a contemporaneous relative regional house price effect, implying that some within-year transmission from London is already factored in.

An important hypothesis concerns the question of stock and flow equilibrium effects on house price determination. The stock equilibrium effect enters through the log income per house variables, $lrynhs_r$ and $lrynhs_{GB}$, discussed above. A flow equilibrium can be examined through the effects of housing stock changes and population changes. The idea is that short term increases in the housing stock relative to population lead to short-term local excess supply, with downward pressure on local prices. Conceivably, this could also reflect an expectations effect in that market participants may believe that a higher rate of house building relative to population growth could have an impact on future house price changes. We measure the effect by including $\Delta \ln(wpop_r/hs_{r,-1})$ in each region's equation. We find a significant effect, suggesting that a 1 percent rise in working age population relative to the housing stock has a short run effect of the order of 1½ to 2 percent on the region's house price index.

[15] The downside risk effect is different from the loss aversion effect set out in Genesove and Mayer (2001), who show that aversion to current nominal losses is important in explaining the positive correlation between house prices and housing transactions.

We also investigated whether the change in the regional proportion of households in the main ages for first time buyers (20 to 39) had any effect. The estimated effect of this variable, pop2039, is statistically significant and positive and enters as a one year lag.

Income dynamics also turn out to be important. Outside London and the South East, we impose the same coefficient on the current year rate of growth of per capita, national disposable non-property income. In London and the South East the income growth coefficients are somewhat higher. The previous year's income growth rate is also important. The region-specific growth rates have little explanatory power, a surprising result. The negative interaction of the credit conditions index, cci, with the current income growth rate suggests the current cash flow constraint matters less as credit becomes more easily available. Finally, there are no explicit income expectation terms in the model, though expectations effects are likely to be reflected in the interest rate, income dynamics and perhaps house price dynamics included in the model.

It is often thought that the behaviour of the stock market, or financial wealth more generally, has an effect on the housing market. We failed to find a positive levels effect from either, unlike earlier national studies by Meen (1996) and Muellbauer and Murphy (1997). This may be because we do not have good regional wealth data. However, the rate of growth of the FTSE index of equity prices in real terms has significant positive effects, in Greater London and the South. This probably reflects the concentration of well-paid employees in the financial services sector in London and the surrounding area, and the link between bonuses, M&A activity and the rise in the stock market.

It is sometimes suggested that relative returns or relative risks in housing and shares influence the allocation of investment flows between the two sectors; for example, that a fall in the stock market leads to a switch into housing due to a rise in the perceived risk of stock returns.[16] A simple measure of downside risk for the stock

[16] It is interesting that the common expression "as safe as houses", meaning a safe or sure bet, was first recorded in 1857 when the railway bubble began to burst and investors turned to speculating in housing.

market can be defined by Δlrftseneg which is equal to the log change in the real FTSE index, Δlrftse, if this is negative and is otherwise zero. This effect is important in Greater London and the South, where share ownership and active portfolio investors are most likely to be concentrated, but irrelevant outside these regions. The two stock market effects together suggest that, for example, a 20% stock market downturn has a much smaller (absolute) effect on house price inflation in Great London and the South than a 20% upturn.[17] This, of course, is in addition to any monetary policy changes that might accompany swings in the stock-market (see Cecchetti et al., 2000). We can accept the hypothesis that the two FTSE effects are zero in the rest of the country.

The regional equations generally include dummy variables for 1988, 1989 and 2001. The year 1988 is special because it became clear that domestic rates (local property tax) would be abolished in England and Wales and replaced by the Poll Tax. It also marked the March announcement that from August 1st, tax relief for mortgage interest would be restricted to one per property. This led to a surge of purchases financed by joint mortgages to meet the August deadline, pushing up prices, especially in London, in the early part of the year. We can accept the restriction that the 1988 and 1989 effects were the same outside London. The 1989 effect was negative in the case of London, probably because of the advancement effect of the tax relief change, which shifted demand from 1989 into 1988.[18] By 1990, it had become apparent in the face of huge collection costs and political protests, that the Poll Tax would not survive and property taxes would return. This rationalises the lack of dummies for 1990 and beyond. However, it is also plain that dummies for 1988 and 1989 could also reflect a house price bubble. We have no simple way of distinguishing between these hypotheses.[19] A dummy for 2001 can be argued to reflect a 9/11 effect, marginally more severe in London, though we can accept the hypothesis of a uniform effect across regions.

[17] The estimated short run, impact effects are a fall of about 1½% and a rise of about 6½% for a 20% downturn / upturn in the real FTSE.

[18] For London, the effects of the restriction of interest tax relief to one per property, rather than one per person sharing a mortgage would have had the sharpest effects given high house prices and a high proportion of young earners.

[19] We also investigated property tax effects (domestic rates and their abolition and subsequent replacement by Council Tax) since variations in tax rates over time and over regions should have effects on prices. Despite pain-staking work constructing regional tax data back to 1975, the estimated tax effects were insignificant. The use of 1988 and 1989 dummies probably picks up much of the effects of the tax switches of the time, as discussed. Handling expectations of tax changes will always be difficult.

--- Table 3 About Here ---

Some single equation diagnostics are presented in Table 3. The first two columns are for the model with a constant speed of adjustment. Overall the model fits well, although there is evidence of mild residual autocorrelation in one equation. The stability of the model was checked by estimating it over different sub-samples and estimates for 1972-1996 are shown.

4. Further Discussion of the Results

In order to get a feel for the magnitude of the various effects in the model, it is useful to look at some figures. Figure 3 shows the estimated long-run effects of the credit conditions index (cci), real and nominal mortgage rates interacted with cci.[20] Note that these merely the products of estimated coefficients and explanatory variables and not a variance decomposition or stochastic simulation. Relative to the 1970s, the estimated effects of cci, in terms of its direct, positive effect on real house prices, is roughly canceled out by the effect of the rise in real interest rates. The figure explains why controlling for cci is crucial to finding the negative effect of real interest rates on house prices predicted by theory: omission of cci would induce serious omitted variable bias. Interestingly, relatively to 2000, the estimated long-run effect of lower interest rates in 2003 is about 18%, but this falls to around 13% when measured against the average interest rate for the period 1985 to 2000.

--- Figures 3 and 4 About Here ---

Figure 4 shows the effects of downside risk, clearly a lagged endogenous variable.[21] It suggests that the depth of the early 1990s housing market recession had much to do with the negative rates of return (and probably the associated payment difficulties and possessions problems faced by homeowners). This was so especially in Greater London, where the effect only began to lift after 1995. Figure 4 also displays the

[20] In terms of the model set out in Table 1, the two effects are βcci * MAcci and (1 – φ * MAcci) * $\beta labmr$ * (demeaned $labmr_{-1}$) + $\beta crabmr$ * MAcci * (demeaned rabmr) + $\beta rabmr$ * (rabmr + sdr) respectively, where MAcci = ½(cci + cci_{-1}). The interest rate / cci interaction reduces the weight on nominal mortgage rates but increases the weight on real rates as cci rises, with an effectively zero weight on the real rate before 1980.

[21] The downside risk effect is $\beta rrhneg_{-1}$ * ($rrhneg_{r,-1}$ + $rrhneg_{r,-2}$ + $rrh.neg_{r,-3}$ + $rrhneg_{r,-4}$).

fitted (asymmetric) effect of the rate of change in the real FTSE share price index, significant in Greater London and the South.

--- Figure 5 About Here ---

Figure 5 shows the effect of changes in the proportion of the working age population aged 20-39, interpreted as a long run effect.[22] This could be defended in terms of the approximate I(1) nature of the data. Interestingly, it plays a considerable role in explaining the out-performance of Greater London house prices in the late 1990s and early 2000s. It also helps explain why house prices were apparently slow to respond to the interest rate rises of 1988-90 - the changing age structure was still supporting the market – as well the weak market conditions between 1992 and 1997.

--- Figure 6 About Here ---

Figure 6 shows composite income, population and housing stock effects.[23] The latter include both the effect of average income per house and the effect of the rate of growth of working age population relative to the lagged housing stock. This suggests that, before 1997 or so, the rate of house building broadly matched rises in real incomes and working age populations (and implicitly household formation). However, since then, the latter have greatly outpaced the rate of house building, especially in Greater London, so driving up real house prices. In Greater London, this was the result both of higher per capita income growth and of population growth, driven by net foreign immigration. However, since 2002 or so, the net change in population has altered, with net outflows from Greater London to other regions offsetting immigration. The overall consequence is that this composite effect explains most of the rise in real house prices since around 1997, thus confirming the relevance of the Barker Inquiry on Housing Supply (Barker, 2004).

--- Figure 7 About Here ---

[22] The demographic effect is $\beta\Delta pop2039_{-1}$ * Δpop2039$_{r,-1}$ / ($\lambda + \lambda_s$.sdr).

[23] The composite income, population and housing stock effect equals θ * ((1 - w_0) * lrynhs$_r$ + w_0 * lrynhs$_{GB}$)) + $\beta\Delta lwpophs$ * Δlog(wpop$_r$/hs$_{r,-1}$) / ($\lambda + \lambda_s$.sdr).

Figure 7 shows one version of an equilibrium correction term including income per house, Greater London catch up, credit and interest rate effects.[24] The change in age structure and the rate of change in population per house, two near I(1) variables in our data, are excluded from this figure. Figure 7 suggests that, given interest rates, incomes, population and housing stock, Greater London was only moderately overvalued in 2003, while the West Midlands and the North were substantially undervalued.

To obtain more causal interpretations from the model, we ran a range of stochastic simulations of counterfactuals for 1998-2003, the main period of appreciation. For example, by holding income per house fixed at the 1998 level, we can estimate the contribution of rising income per house – a mix of higher income per head, population growth and lower rates of house-building. We discover that weighted average of regional house prices would have been 25 percent lower in 2003 with the counterfactual. An alternative counterfactual in which nominal interest rates are held at their 1998 values until 2003, results in real prices being 7.5 percent lower. This will understate the interest rate effect somewhat, since part of it works through real income and through equity prices. The other counterfactuals, keeping cci, inflation, the change in the proportion of the population aged 20 to 39, and the Stamp Duty rate all at their 1998 values and setting to zero growth in the real FTSE index, have a composite effect of under 2 percent, with single effects all in the range of 0.7 to -1.8 percent.

5. Was There a House Price Bubble?

In this Section, we address the question raised in the Introduction of whether or not there has recently been a bubble in UK house prices.[25] Much of the debate about house price bubbles focuses on the time series behaviour of ratios of house prices to income, or mortgage payment to income ratios. These ratios are not very informative

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[24] The error correction term in Figure 7 equals:

$$\mathrm{lrhp}_{r,-1} - \beta 0_r - \beta year_r * (\mathrm{year} - 1990) - \theta * ((1 - w_0) * \mathrm{lrynhs}_r + w_0 * \mathrm{lrynhs}_{GB}))$$
$$- \beta cci * \mathrm{MAcci} - (1 - \varphi * \mathrm{MAcci}) * \beta labmr * (\text{demeaned labmr}_{-1})$$
$$- \beta crabmr * \mathrm{MAcci} * (\text{demeaned rabmr}) - \beta rabmr * (\mathrm{rabmr} + \mathrm{sdr}).$$

where, as before, $\mathrm{MAcci} = ½(\mathrm{cci} + \mathrm{cci}_{-1})$.

[25] Inter alia, Farlow (2004) summarizes, in a non-technical manner, many of the factors that could give rise to a house price bubble. Meen (2000) shows that, at the national level, persistent excess returns can be earned on housing so the UK market is inefficient. At a local level, Rosenthal (2004) finds that monthly house prices are not very predictable.

about the presence of absence of bubbles, because they ignore a range of other important factors. For example, as we have shown, demographics and new housing supply effects are important in the short run as well as the long run.[26]

Another strand of the housing bubbles literature looks at the ratio of house prices to rents, using an equilibrium asset pricing approach. For example, see Ayuso and Restroy (2003) and Weeken (2003). This approach appears very attractive and simple, since house prices as such do not have to be modelled. The basic problem with this approach is the small size of the private rented sector in the UK, which is not representative of the private housing sector as a whole, and the poor quality of the available rent data. Furthermore, even with good data on rents, demand shocks will shift price to rent ratios because rents are far stickier than house prices. Moreover, a number of auxiliary assumptions are required to implement this approach, as acknowledged by both Ayuso and Restroy (2003) and Weeken (2003). Nevertheless, Ayuso and Restroy (2003) suggest that UK house price to rent ratios were about 20% above their equilibrium value in 2002. Weeken (2003), whose results imply that house price to rent ratios were only a few percentage points above their equilibrium level in 2002, suggests that "because of data and model limitations, no firm conclusions can be drawn".

A third strand of the housing bubbles literature is more technical. For example, Roche (2001) estimates a regime switching model for Dublin house prices. Special cases of this model are a fads model and a partial collapsing (speculative) bubble model.[27] The regime switching model is estimated in two stages. In the first stage, the non-fundamental component of house prices is estimated. In the second stage, the actual regime switching model is estimated using last period's estimated non-fundamental prices as the only explanatory variable explaining the change in house prices this period. This means that the regime switching model results are crucially dependent on the model used to estimate the fundamental and non-fundamental components of house prices so modeling the determinants of house price cannot be

[26] The debate in the US pays more attention to these factors. See Case and Shiller (2003) and Himmelberg, Meyer and Sinai (2005), inter alia.

[27] In both the fads and collapsing bubbles models, houses prices may systematically differ from fundamentals over a number of years. In the fads model, the non-fundamental component of house prices is mean reverting. However, in the collapsing bubbles market, there is a period when the non-fundamental or speculative component of house prices grows along with the probability of a collapse in this component.

avoided. Moreover, regime switching models can be difficult to estimate since the models are highly non-linear. In practice, long runs of high frequency data are required which often means that a limited set of explanatory variables are used.

We have not followed the switching models of Zhou and Sornette (2003) and Garino and Sarno (2004) since they also lack fundamentals. The former only look at house price data to search for signs of faster than exponential growth and periodic oscillations in log house prices, signalling an unsustainable regime and endogenous bubbles respectively. Garino and Sarno (2004) use Markov-switching unit root tests and a test for cointegration which is robust to the skewness and excess kurtosis encountered when there are periodically collapsing bubbles. Unfortunately, their model of fundamental house prices is lacking; house prices only depend on incomes and interest rates.

--- Table 4 About Here ---

Implicit in a final strand of the housing bubbles literature is that the market is inefficient, for example, because of the lack of investors with 'deep pockets' (Stein 1995). We estimate a dynamic equilibrium correction model. As Abraham and Hendershott (1996) have argued, the lagged house price growth term in such a model can be given a 'bubble builder' interpretation, while the lagged equilibrium correction levels effect reflect the 'bubble burster' influence of prices departing too much from fundamentals. This is the approach which we and many others take (Bourassa et al. 2001 and Capozza et al. 2004). For example, the IMF (2003,2005) estimate a dynamic, reduced form error correction equation for log real house prices as a function of log real disposable income per household and real interest rates only. The resulting equation is not really a reduced form since no supply side variables such as construction costs are included. Unfortunately, it is also rather unstable. This is not surprising since, supply, the changing age structure of the population, nominal interest rates or shifts in UK credit conditions play no role.

Barrell et al. (2004) also estimate a dynamic, error correction equation relating log real house prices to log real disposable income, real interest rates, (in the short run) nominal interest rates and dummy variables. They could not obtain a significant housing stock effect. Both Barrell et al. (2004) and the IMF (2005) suggest that, on the basis of their model simulations, UK house prices were overvalued by 30% or more in 2003.

We believe that the IMF (2003, 2005) and Barrell et al. (2004) house price equations are mis-specified and that our more general model, which has a sound theoretical basis and is data consistent and stable over time, is better for testing the bubbles hypothesis.[28]

Our model results do not suggest that house prices in 2003 are substantially overvalued. Fitting our model to data for 1972 to 1996 to forecast 1997-2003 gives no signs of systematic under-prediction either for the full period or for 2000 to 2003. When the regional forecast errors are aggregated up to the level of Great Britain, the mean and mean absolute forecast errors are about -0.5% and 3.2% respectively for the period 1997 to 2003, though for 2003 the price level is under-predicted by 4 percent.[29] For the full sample, our model fits the data well, with little left over to be classified as bubble. Moreover, the in-sample forecasts suggest that we have not over-fitted the model, which might have compromised bubble detection.

We have also examined out-of-sample scenarios to see what the model suggests about the course of real house prices in the next few years. If the model were to suggest that house prices might collapse in some circumstances, then house prices could arguably be overvalued by 20% or more. Inter alia, we considered the two scenarios set out in Table 4.

--- Figure 8 About Here ---

The base scenario, which involves a mild slowdown in the economy for two years, is a little more pessimistic than the consensus forecasts made in March 2006. The simulation results in Figure 8 provide no evidence of a house bubble. We explored the robustness of this result by simulating the effects of a 50% increase in the rate of growth of the housing stock (albeit from a very low level). We found that real house

[28] The model of house price speculation in Levin and Wright (1997a, 1997b) is interesting but is not consistent with the standard or textbook model of house prices set out in Section 2.

[29] In the forecasts, the 2001 dummy variable effect is treated as a genuine shock. Excluding 2001, the mean and mean absolute forecast errors are 0.2% and 2.9%.

price growth would only be marginally lower, though the effect on the level of real house prices accumulates over time.

We also considered an improbably "gloomy" scenario in which the economy turns sour – inflation rises, interest rates rise in response, per capita real income falls slightly and the stock market is flat for two years before real values grow again. In this case, assuming there are no REITS valuation effects[30], the simulation implies that moderate nominal falls in house prices in 2006-2007 are a possibility, especially in London and the South[31] but not in the country as a whole. Again, the results suggest that there is no bubble in house prices.

6. Conclusions

The regional house model presented here has the necessary features of a valid model suggested by Meen and Andrews (1998) and Meen (2001). The model is data consistent; incorporates spatial lags and errors; has some spatial coefficient heterogeneity; has a plausible long run solution; includes a full range of explanatory variables.[32]

Our model captures long run fundamentals such as the effect of income, population, age composition, the housing stock, and interest rates on the long-run level of real house prices. It also builds in the effect of house price dynamics, including transmission from leading or adjacent regions on other regions. In the UK, the leading region is London, and the "ripple effect" of changes there impacting on other regions, first on adjacent ones, is a notable feature in the UK.

We distinguish between the short-run and long-run effects of house-building and population growth. We allow for stock market effects and test for heterogeneity

[30] Legislation easing the tax regime for Real Estate Investment Trusts was under discussion in 2004 - 2006. This might be expected to increase investment in rental housing.

[31] If the consumption, income and exchange rate feedbacks are large, the fall in nominal house prices could be self-reinforcing resulting in a larger downturn, operating partly through the downside risk effect. This downturn would be temporary if the global interest rate environment remains kind, given the high levels of debt and house prices in the UK and US.

[32] . When using the model for forecasting or simulating policy scenarios, feedbacks to and from regional house prices, earnings, employment and unemployment and inter-regional migration, as well as more conventional and important macro feedbacks from interest rates etc. should be factored in. See Cameron and Muellbauer (1998, 2001), Cameron, Muellbauer and Murphy (2005) and Johnes and Hyclak (1994, 1999) inter alia.

between regions in some of the key parameters. We also examine the effect of easier credit conditions resulting from structural changes in UK credit markets. Easier credit conditions not only have a direct effect on the level of real house prices, but also shift the relative roles of real and nominal interest rates: the former become more important and the latter less important. This links the nominal rate effect with 'tilt' or 'front-end loading' effect, which declines as access to credit improves. Our findings suggest Brunnermeier and Julliard (2006) are incorrect to eliminate the tilt interpretation of the nominal interest rate effect on house prices.

The great advantage of our regional panel model is more precise estimation of the structural parameters, which makes it possible to draw more robust conclusions. A striking finding is that estimating the model on data up to 1996 and forecasting conditionally on the other variables leads to no symptoms of systematic under-prediction of house prices in the period 1997-2003 when house prices rose very strongly. Furthermore, our conclusion is that the evolution of regional prices in this period can be largely explained by the combination of strong income growth, higher population growth (partly from in-migration), lower interest rates and low rates of house-building. However, for 1988-9, we cannot rule out the possibility that the significant dummies reflect an element of 'bubble' behaviour as well as a rational response to the tax changes of the time.

Finally, as a further check, we simulated house price developments for the period 2004-2010 on a range of assumptions about income growth, population growth, house-building, inflation and interest rates. We found that only quite negative scenarios – more negative than any currently contemplated by main-stream forecasters - would produce falls in nominal house prices. London and the South are the regions where such negative scenarios would have the largest regional effects, if they occurred. The question, 'was there a recent house price bubble in the UK' thus appears to have the answer 'no'. It would be paradoxical if one of the factors preventing a house price bubble was the publicity given to those warning of the risk of the bubble collapsing.

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Table 1
House Price Equation for
Regions other than Greater London, the South and Scotland

$\Delta\text{lrhp}_r =$	Dependent variable = change in log real house prices in region r, adjusted for changes in mix and survey coverage.
$(\lambda + \lambda_s\text{*sdr}) * \beta 0_r$	Region specific intercept scaled by speed of adjustment varying with the rate of Stamp Duty sdr. Speed of adjustment scales all long-run effects.
$+ (\lambda + \lambda_s\text{*sdr}) * \beta year_r * (\text{year - 1990})$	Region specific linear trend.
$+ (\lambda + \lambda_s\text{*sdr}) *$ $[\theta*((1 - w_0) * \text{lrynhs}_r + w_0 * \text{lrynhs}_{GB}) - \text{lrhp}_{r,-1}]$	Equilibrium correction term. The long run elasticity of house prices w.r.t. real (non-property personal disposable) income and the housing stock θ is 1.6. Regional and national income per house are weighted by 1- w_0, and w_0, which equals 0.67. The values of θ and w_0 are both data admissible.
$+ (\lambda + \lambda_s\text{*sdr}) * \beta cci * \text{MAcci}$	Positive effect of average credit conditions. The cci measure is from Fernandez-Corugedo and Muellbauer (2005). MAcci = ½(cci + cci_{-1}).
$+ (\lambda + \lambda_s\text{*sdr}) * (1 - \varphi * \text{MAcci}) *$ $(\beta\Delta_2 labmr * \Delta_2\text{labmr} + \beta labmr * \text{demeaned labmr}_{-1})$	Negative effect of two period change in log (tax adjusted) mortgage rate and lagged level of the same variable interacted with the credit conditions measure cci. Nominal interest rates matter less when credit is more freely available.
$+ (\lambda + \lambda_s\text{*sdr}) * \beta crabmr * \text{MAcci * demeaned rabmr}$	Negative effect of real tax adjusted mortgage interest rate. Real interest rates matter more as credit becomes more freely available.
$+ (\lambda + \lambda_s\text{*sdr}) * \beta rabmr * (\text{rabmr + sdr})$	Negative effect of sum of real tax adjusted mortgage interest rate plus Stamp Duty rate.
$+ \beta\Delta rlhp_{-1} * [(1 - w_{1r} - w_{2r}) * \Delta\text{lrhp}_{r,-1} + w_{1r} * \Delta\text{clrhp}_{r,-1}$ $+ w_{2r} * \Delta\text{lrhp}_{GL,-1}]$	Positive effect of lagged change in real house prices in the region, in contiguous regions and in Greater London. The weights are region specific.
$+ \beta\Delta lrpdin * \Delta\text{lrpdin}$	Positive effect of changes in national non-property income (pdin).
$+ \beta\Delta lrpdin_{-1} * \Delta\text{lrpdin}_{-1}$	Positive effect of lagged change in non-property income.
$+ \beta cci\Delta lrpdin * \text{MAcci} * \Delta\text{lrpdin}$	Negative interaction of cci and Δlrpdin. Households are less cash constrained if cci is high, so income changes matter less.
$+ \beta\Delta^2 lpc * \Delta^2\text{lpc}$	Negative effect of acceleration in the inflation rate.
$+ \beta rrhneg_{-1} * (\text{rrhneg}_{r,-1} + \text{rrhneg}_{r,-2} + \text{rrh.neg}_{r,-3} + \text{rrhneg}_{r,-4})$	Negative downside risk effect, using MA_4 of lagged negative real rates of

	return on housing.
$+ \beta\Delta pop2039_{-1} * \Delta \text{pop2039}_{r,-1}$	Positive demographic effect captured by changes in the share of the working age population aged 20 to 39 in the region.
$+ \beta\Delta lwpophs * \Delta \ln(\text{wpop}_r/\text{hs}_{r,-1})$	Positive effect of change in the ratio of the working age population to the housing stock in each region.
$+ \beta 88 * \text{d88}$ $+ \beta 89 * \text{d89}$	Positive time dummies for 1988 and 1989. Poll Tax and advance announcement of end of multiple mortgage interest tax relief.
$+ \beta 01 * \text{d01}$	Negative time dummy for 2001 capturing 9/11 and stock market turmoil effects.

Notes: MAcci = ½(cci + cci_{-1}). The Greater London and South equations are similar to the equation set out above except that (i) some of the short run coefficients and time dummies are allowed to take different values and (ii) positive and negative changes in the real value of the FTSE are included as additional explanatory variables ($\beta\Delta lrftse * \Delta\text{lrftse} + \beta\Delta lrftse\ neg * \Delta\text{lrftse neg}$). Thus, changes in the real value of the FTSE have an asymmetric effect in GL and in ST: the negative effect of a fall in the real FTSE is much smaller that the positive effect of a rise in the real FTSE. The Scottish equation is the same as set out above except that the 1988 time dummy is zero.

Table 2
Two-Step Model Parameter Estimates
Dependent Variables = $\Delta lrhp_r$

Variables / Composite Variables		Basic Model				Variable Adjustment Model			
		1972-2003		1972-1996		1972-2003		1972-1996	
		Coeff	t Stat	Coeff	t Stat	Coeff	t Stat	Coeff	t Stat
Speed of Adjustment	λ	0.285	14.4	0.354	12.3	0.372	8.9	0.422	5.7
	λ_s (Stamp Duty)	0	-	0	-	-7.183	-3.7	-8.152	-1.2
Lagged House Price Growth		0.460	18.8	0.481	19.7	0.499	16.1	0.500	11.7
Credit Conditions MAcci = ½(cci+cci_{-1})		1.021	4.3	0.506	2.3	1.131	3.9	0.624	2.8
Income Growth With Credit Effects	Δlrpdin (ex. GL,ST)	0.836	8.9	0.753	7.1	0.857	9.8	0.770	8.6
	Δlrpdin (GL)	1.015	15.8	0.857	10.4	0.996	12.4	0.886	11.9
	Δlrpdin (ST)	0.897	7.2	0.828	6.4	0.901	7.3	0.843	6.9
	MAcci * Δlrpdin	-3.397	-5.2	-4.791	-6.9	-4.028	-6.0	-4.943	-6.9
	$\Delta lrpdin_{-1}$ (ex. GL,ST)	0.473	6.6	0.487	7.0	0.467	7.0	0.503	6.8
	$\Delta lrpdin_{-1}$ (GL)	0.847	7.9	0.951	9.4	0.804	7.5	0.879	7.9
	$\Delta lrpdin_{-1}$ (ST)	0.473	3.7	0.515	3.9	0.453	3.5	0.495	3.5
Interest Rate With Credit Effects	(1 – φ.MAcci) * $\Delta_2$labmr (ex. GL,ST)	-0.273	-4.1	-0.273	-4.6	-0.242	-3.8	-0.276	-4.6
	(1 – φ.MAcci) * $\Delta_2$labmr (GL)	-0.343	-3.5	-0.333	-3.6	-0.328	-3.3	-0.331	-3.2
	(1 – φ.MAcci) * $\Delta_2$labmr (ST)	-0.345	-4.2	-0.333	-4.6	-0.314	-3.6	-0.333	-3.9
	(1 – φ.MAcci) * $labmr_{-1}$	-0.348	-6.8	-0.277	-5.9	-0.319	-6.8	-0.275	-5.7
	φ	2.118	4.6	2.988	6.7	1.869	3.3	2.750	5.3
	MAcci * rabmr	-27.762	-5.9	-12.866	-3.6	-22.430	-4.2	-12.228	-3.3
	rabmr + sdr	-	-	-	-	-0.632	-2.4	-0.330	-1.0
Downside Risk MA_4 $rrhneg_{r,-1}$		0.248	7.3	0.282	4.1	0.263	5.3	0.268	3.7
Change in Inflation Δ^2lpc		-0.740	-8.8	-0.762	-10.5	-0.824	-8.3	-0.800	-6.9
Demographic Effect $\Delta pop2039_{r,-1}$		0.615	1.0	0.830	1.1	1.233	2.3	1.114	1.4
New House Effect $\Delta(lwpop_r - lhs_{r,-1})$		1.400	8.2	1.409	7.8	1.735	9.4	1.705	7.7

ΔStock Market (GL & ST)	ΔlrFTSE (GL)	0.214	4.7	0.314	5.9	0.276	5.6	0.348	6.6
	ΔlrFTSE (ST)	0.202	6.4	0.216	5.5	0.209	6.8	0.224	5.8
	ΔlrFTSE if neg (GL)	-0.086	-1.4	-0.190	-2.6	-0.167	-2.5	-0.246	-3.1
	ΔlrFTSE if neg (ST)	-0.188	-4.6	-0.204	-4.1	-0.199	-4.9	-0.218	-4.4
Time Dummies	1988 (ex. SC)	0.095	15.0	0.103	15.0	0.100	16.4	0.106	14.3
	1989 (ex. GL)	0.104	36.3	0.107	26.1	0.105	35.2	0.105	22.9
	1989 (GL)	-0.018	-2.9	-0.016	-3.2	-0.023	-3.8	-0.024	-4.7
	2001	-0.064	-12.3	-	-	-0.060	-11.4	-	-

Notes: (i) Two step parameter estimates - seemingly unrelated regression (SUR) followed by GLS with a fixed covariance matrix from SUR. The full sample variance-covariance matrix is used for the shorter 1972 to 1996 sample. (ii) Robust standard errors are used. (iii) MAcci = ½(cci+cci_{-1}). (iv) Regional intercepts and time trends are not reported. The average estimated regional trend coefficient is 0.018. (v) Two long run restrictions are imposed in the equilibrium correction term. The long run income / housing stock elasticity of house prices θ is set to 1.6 and the weight on the national income and housing stock w_0 is set at 0.67. (vi) In the lagged house price increase term, the weights on own region, contiguous region and Greater London house prices ($\Delta rlhp_{r,-1}$, $\Delta crlhp_{r,-1}$ and $\Delta rlhp_{GL,-1}$) are constrained to lie in the [0,1] interval. The contiguous region and Greater London weights are as follows:

Weights	Region								
	North NT	Yorkshire & Humberside YH	East Midlands EM	West Midlands WM	Greater London GL	South ST	South West SW	Wales WW	Scotland SC
Contiguous Region	0.495	1	0.718	0	0	0	0	1	0
Greater London	0	0	0.112	0.280	1	1	1	0	0
Own Region	0.505	0	0.170	0.720		0	0	0	1

Table 3
Some Single Equation Diagnostics
Variable Adjustment Model 2-Step Results for 1972 to 2003
Dependent Variables = $\Delta lrhp_r$

Region	Dep. Var.		Equation		P Values	
	Mean	Std. Dev.	Std. Error	R^2	LM Hetero	LM Auto AR(2)/ MA(2)
NT - North	0.032	0.084	0.026	0.906	15.9%	9.1%
YH – Yorkshire & Humberside	0.036	0.089	0.026	0.894	0.9%	5.3%
EM - East Midlands	0.040	0.107	0.030	0.927	11.1%	24.0%
WM - West Midlands	0.036	0.102	0.018	0.972	1.8%	10.9%
GL - Greater London	0.041	0.118	0.028	0.940	0.7%	21.1%
ST - South	0.040	0.120	0.029	0.941	68.1%	63.4%
SW - South West	0.040	0.121	0.029	0.947	45.0%	3.1%
WW - Wales	0.035	0.096	0.030	0.897	2.0%	39.2%
SC - Scotland	0.025	0.058	0.021	0.869	21.2%	42.0%

Table 4
Some Scenarios

Scenarios	Growth Rates Etc.	Year						
		2004	2005	2006	2007	2008	2009	2010
Base Scenario	Real non-property income	2.1%	1.5%	1.5%	2.0%	2.3%	2.5%	2.5%
	Inflation rate	1.3%	2.5%	2.7%	2.8%	2.6%	2.4%	2.2%
	Mortgage interest rate	5.0%	5.5%	5.5%	5.5%	5.0%	5.0%	5.0%
	Real FTSE	9.0%	8.0%	7.0%	5.0%	5.0%	5.0%	5.0%
Alternative Gloomy" Scenario	Real non-property income	2.1%	1.2%	0.5%	0.5%	1.0%	1.5%	2.0%
	Inflation rate	1.3%	2.5%	3.0%	2.8%	2.8%	2.6%	2.4%
	Mortgage interest rate	5.0%	5.5%	6.5%	6.0%	5.5%	5.5%	5.0%
	Real FTSE	9.0%	8.0%	0%	5.0%	5.0%	5.0%	5.0%

Notes: The real non-property income growth rates are aggregate and not per capita figures.
Other assumptions: The index of credit conditions cci is assumed unchanged. The annual rate of growth of regional housing stocks and regional proportions of owner occupiers over the period 2004 to 2010 is assumed equal to the average annual rate of growth in the previous seven years. The source of our regional population numbers are the population projections produced by the Government Actuaries Department (national figures) and the Office of National Statistics (sub-national figures). Their projections show a decline in the growth of the working age population over the period 2004 to 2010. They also show a further decline in the proportions aged 20 to 39. The largest decline is around 2006. The decline tails off a little after this.

Figure 1: Log Real House Prices

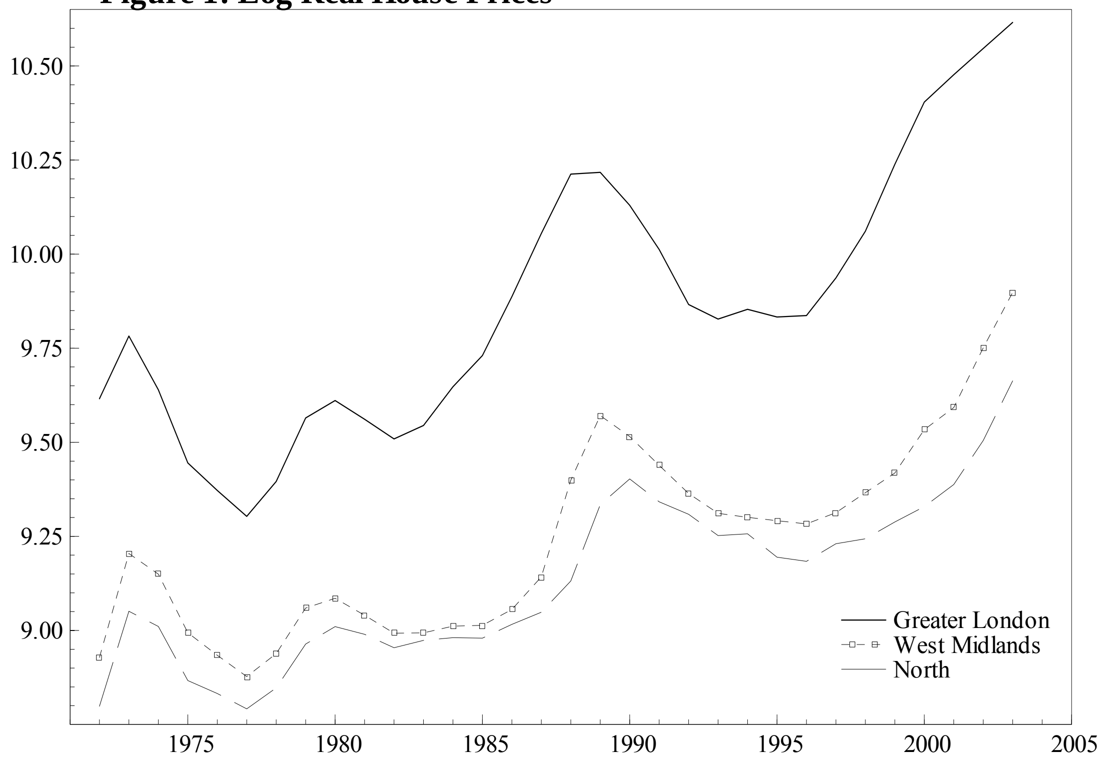
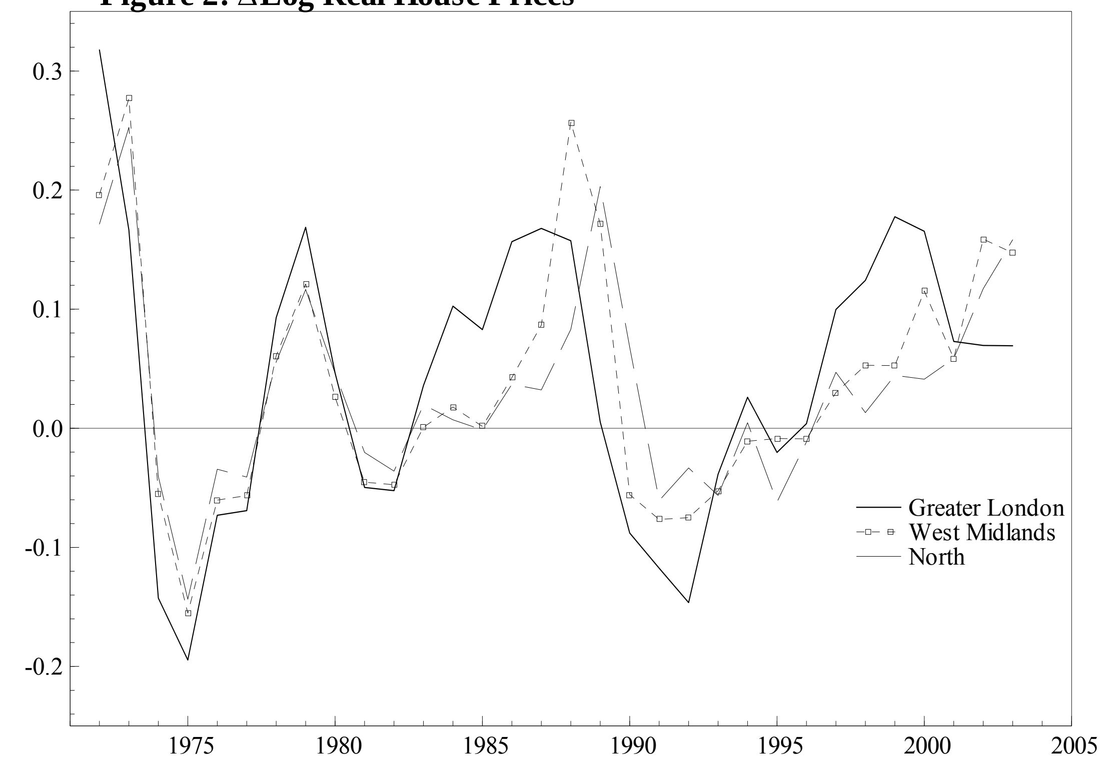


Figure 2: ΔLog Real House Prices



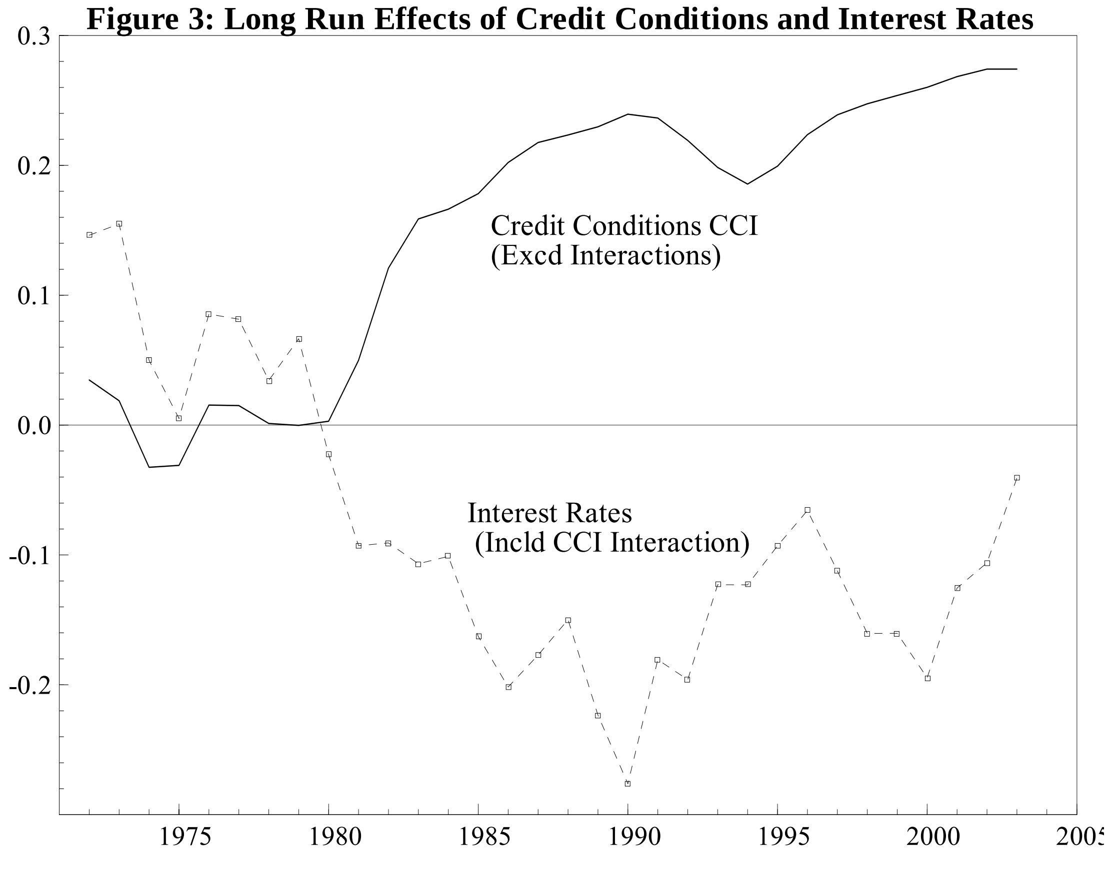


Figure 3: Long Run Effects of Credit Conditions and Interest Rates
0.3
0.2
0.1
0.0
-0.1
-0.2
Credit Conditions CCI
(Excd Interactions)
Interest Rates
(Incld CCI Interaction)
1975
1980
1985
1990
1995
2000

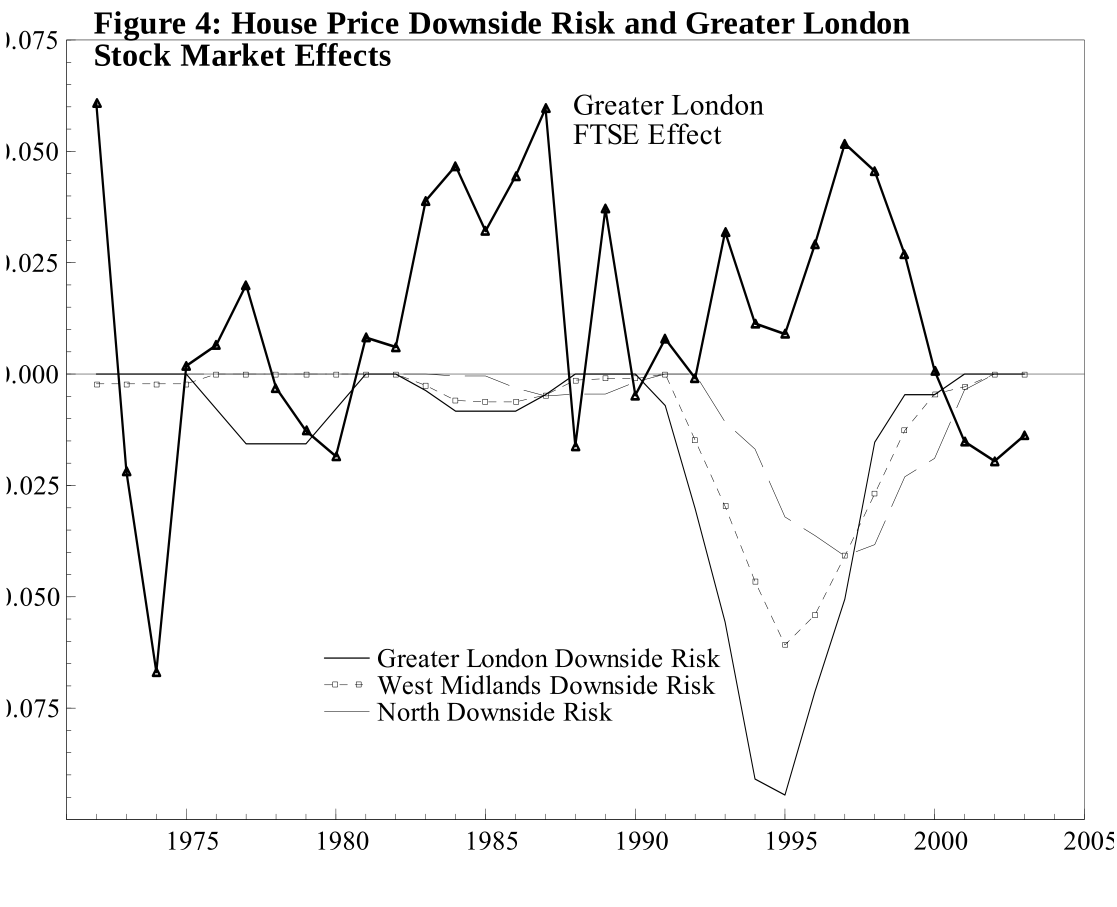


Figure 4: House Price Downside Risk and Greater London
Stock Market Effects
Greater London
FTSE Effect
Greater London Downside Risk
West Midlands Downside Risk
North Downside Risk
1975
1980
1985
1990
1995
2000

Figure 5: Effects of Population Proportion Aged 20 to 39 (in Long Run)

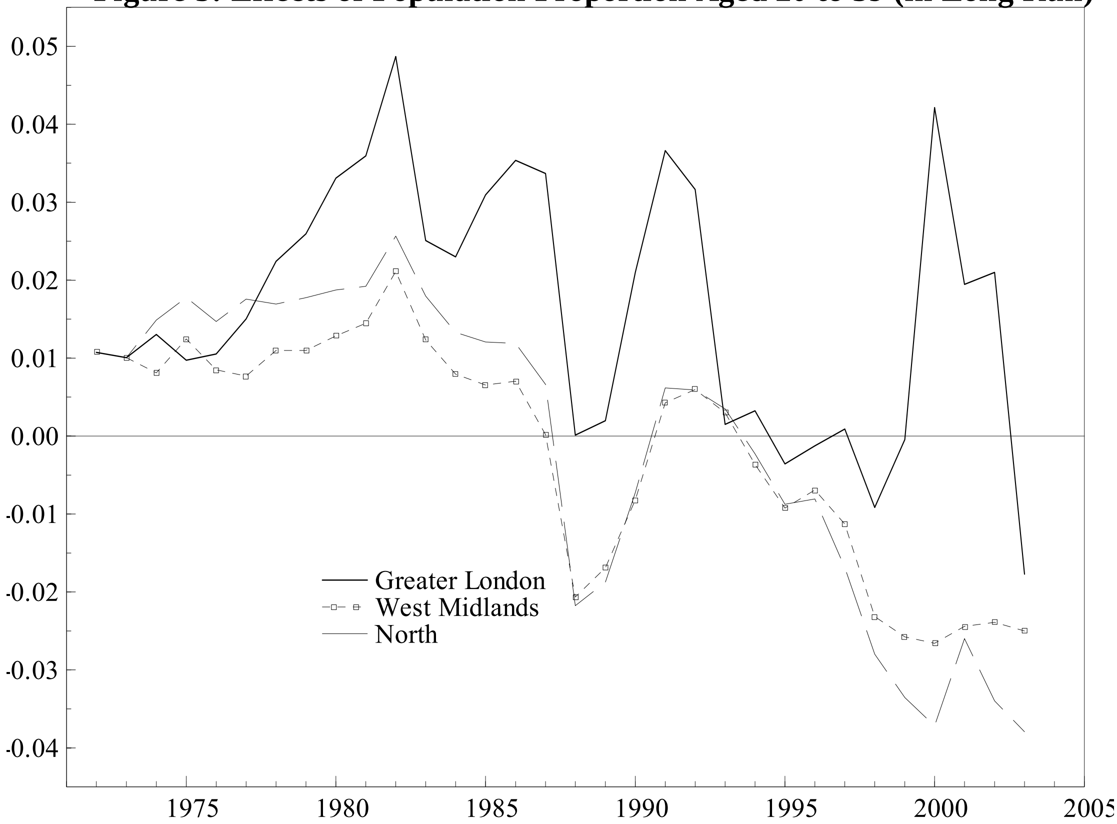


Figure 6: Composite Long Run Income, Population and Housing Stock Effects Incld. ΔLn (WPOP/HS) Effect

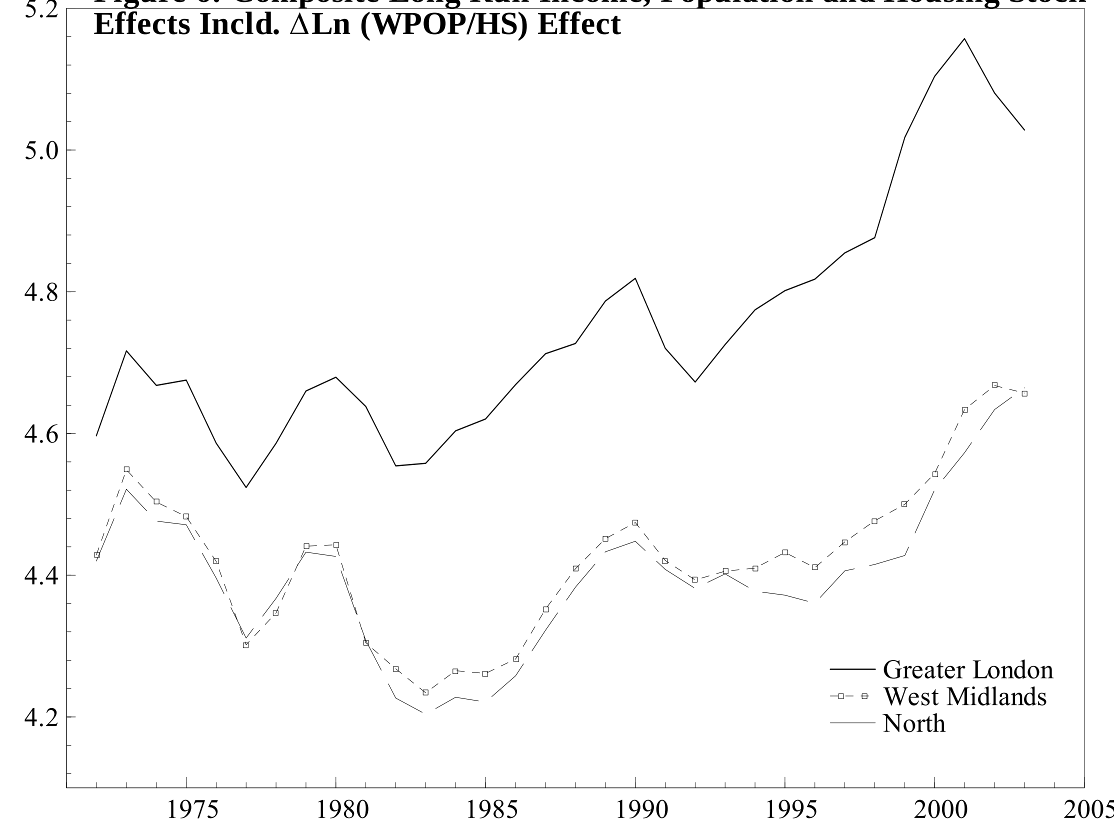


Figure 7: Equilibrium Correction Terms

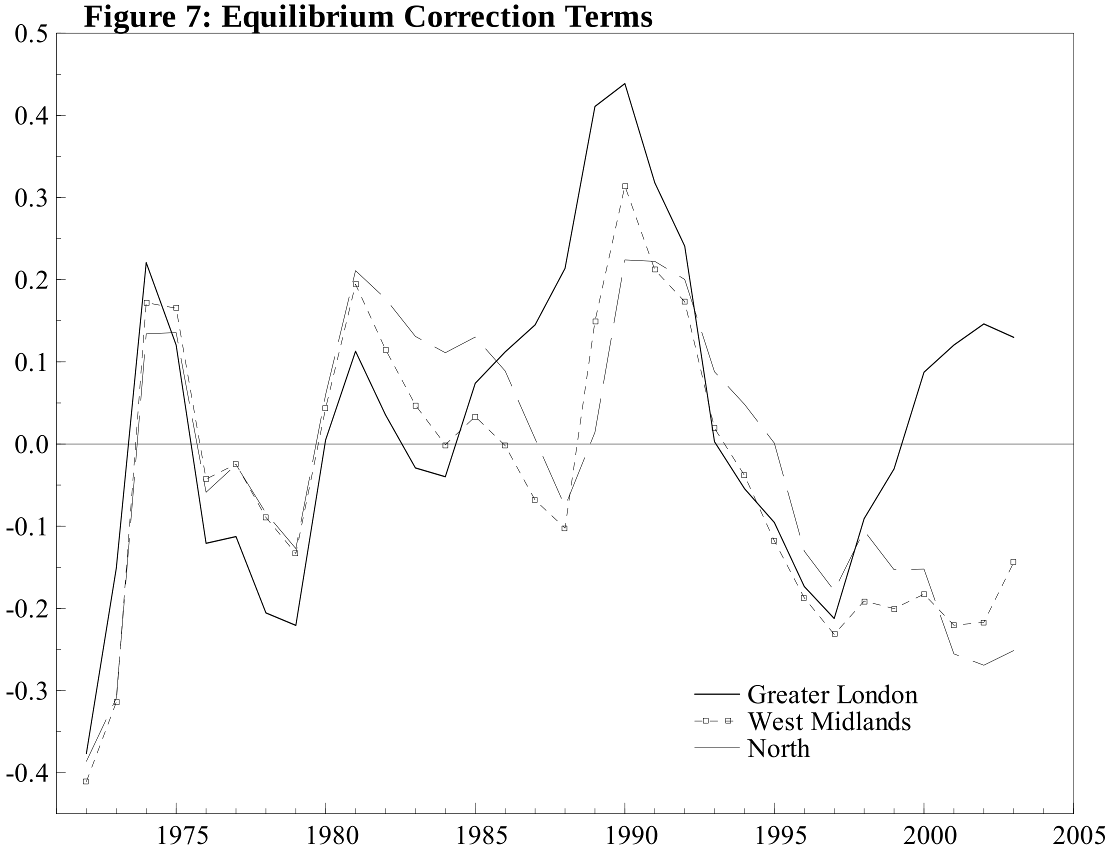
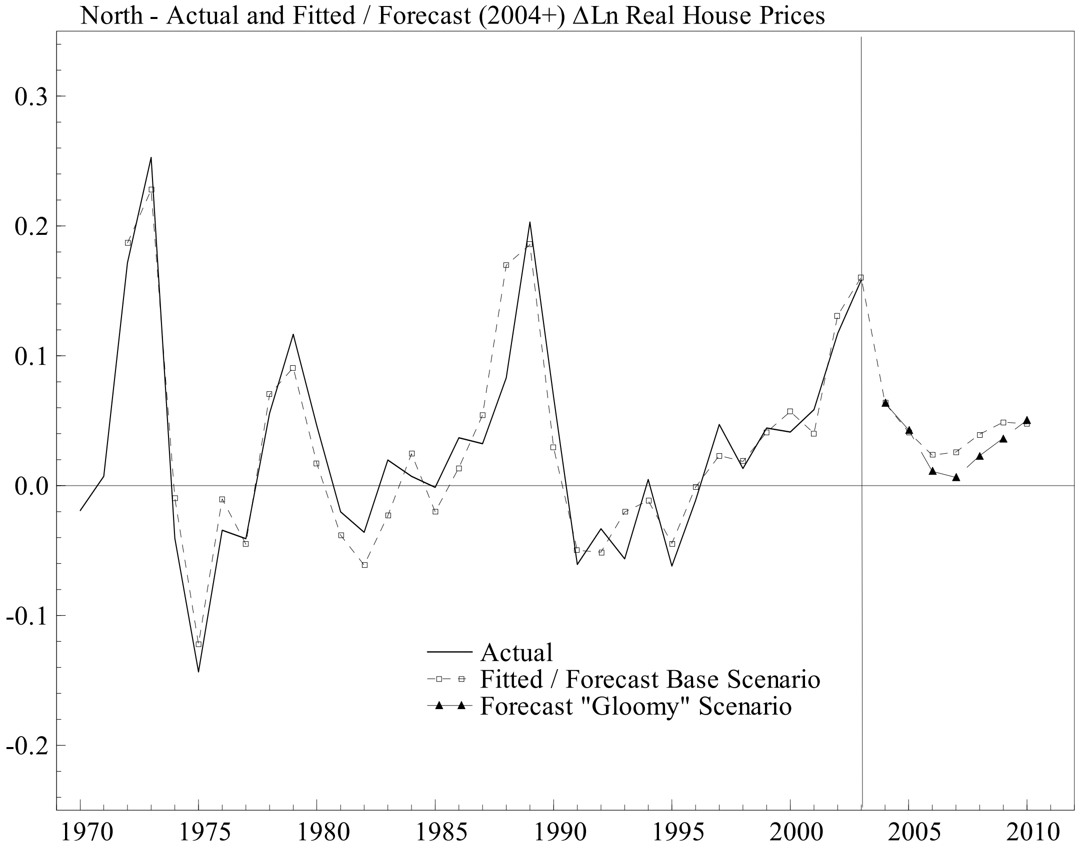


Figure 8 – Out of Sample Simulations



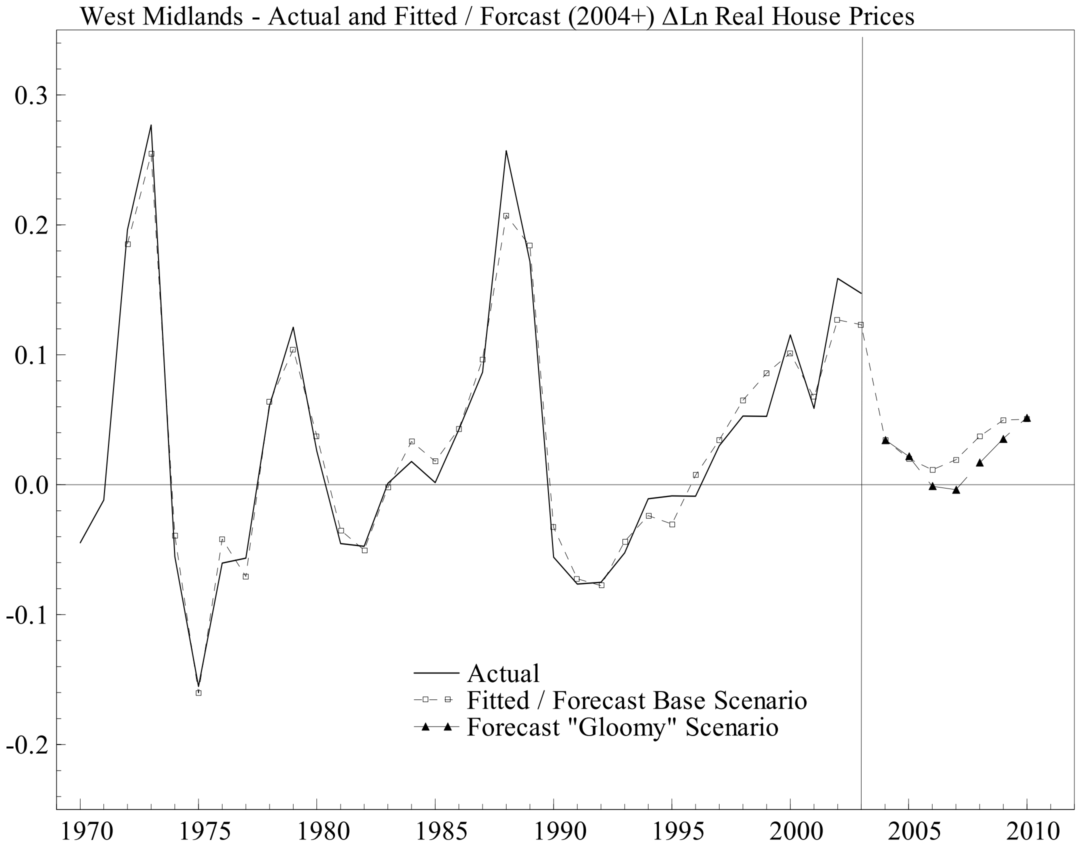
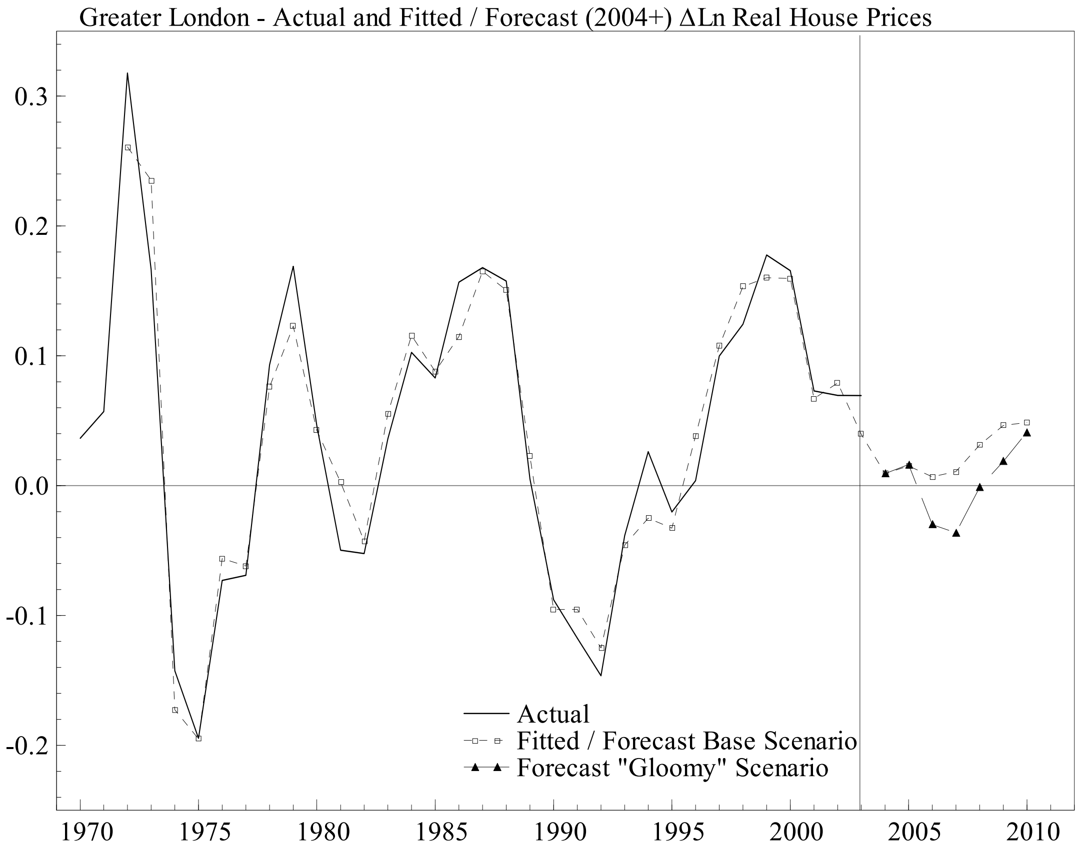
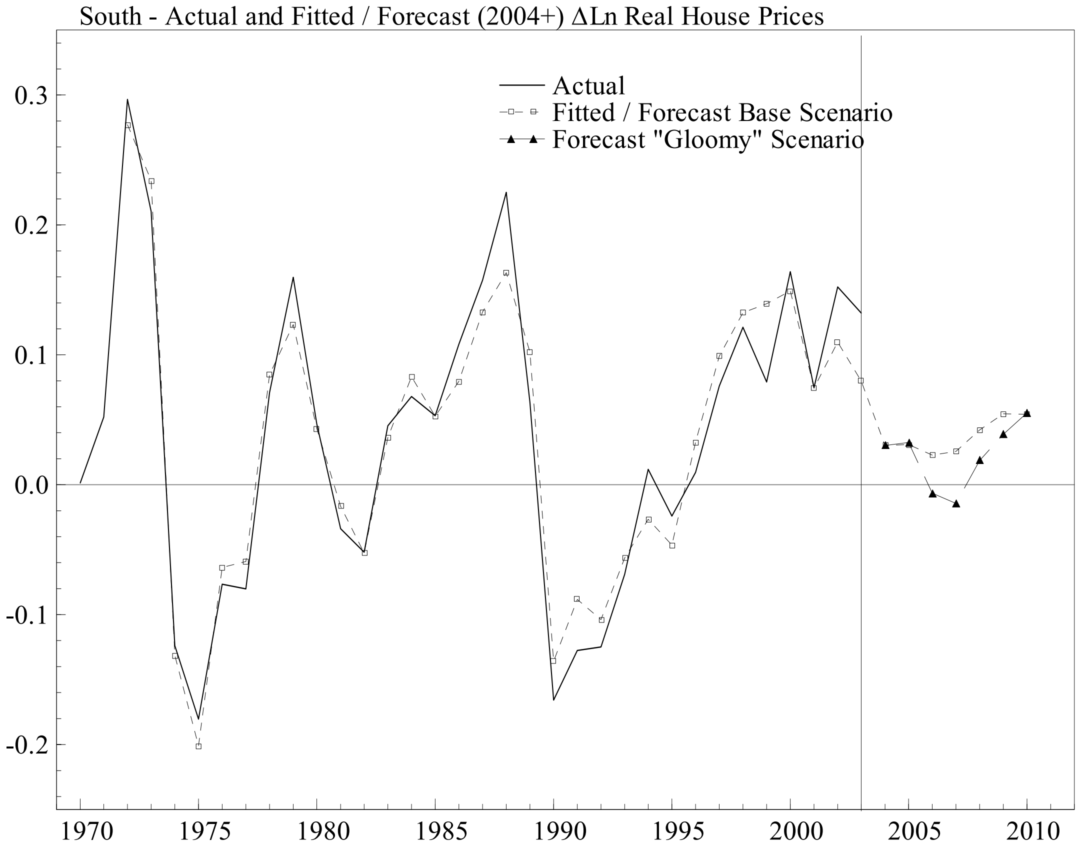


Figure 8 (Continued) – Out of Sample Simulations





Appendix: Data Construction and Sources

(a) House prices:-

All regional and national log house price indices, which are derived by linking official published Office of the Deputy Prime Minister (ODPM) mixed adjusted second hand house price indices, are adjusted by adding an adjustment factor which corrects for composition changes as banks etc. entered the mortgage market (Muellbauer and Murphy, 1997). All indices have been rebased to 1990 using average second hand house price values.

(b) Non-property personal disposable income:-

Log real non-property personal disposable income (pdi) in region r ($rlryn_r$) is defined as follows:

$$rlryn_r = lrpdin + (1/3*rlfte_{r,+1} + 2/3*rlfte_r) + rlempr_r$$
$$+ rlwpop_r + rltaxadj_r + \log((1\text{-}spt_r) + rept_r * spt_r$$

where lrpdin = log real non-property pdi in UK, non-property pdi = (1 - tuk) * (wage and salaries + mixed income) and tuk = 1 - (post tax pdi / pre tax pi). Sources: Office of National Statistics (ONS) Blue Book for wages and salaries (qwlt) and mixed income (qwlm); Oxford Economic Forecasting (OEF) regional model databank for tuk.

$rlfte_r$ = Log relative total full earnings in April in region r, relative to GB. Source: ONS New Earnings Survey (NES) data linked to Annual Survey of Hours and Earnings (ASHE) data from 2002. We used ⅔ of current earnings and ⅓ of next year's earnings because the data are for April.

$rlempr_r$ = Log relative employment rate in region r. Source: OEF regional model data.

$rlwpop_r$ = Log relative working age population in region r. Source: OEF regional model data;

$rltaxadj_r$ = Log relative (post tax pdi / pre tax pi) in region r. Source: OEF regional model data.

spt_r = Share of part time employment in total employment in region r. Source: ONS.

$rept_r$= Ratio of average part time to full time earnings in region r. Source: ONS NES, assumed unchanged post 2001 and pre 1975.

(c) Log income per house:-

$$lrynhs_r = rlryn_r - \log(hs_{r,-1}) - 0.7*\log(poo_{r,-1})$$

where hs_r = housing stock in region r and poo_r = proportion of owner occupiers in region r. Source: ODPM housing statistics.

(d) Return on housing:-

$$rrh_r = \Delta lhp_{r,-1} + 0.03 - abmr$$

$$rrhneg_r = rrh_r * 1(rrh_r<0)$$

where Δlhp_r = first difference of log house price index in region r (source: linked ODPM data); abmr = tax adjusted building society mortgage rate (bmr) with the adjustment based on Inland Revenue estimates of the cost of tax relief (sources: ONS Financial Statistics and HM Revenue & Customs Inland Revenue Statistics) and $1(rrh_r < 0) = 1$ if rrh_r is negative and 0 otherwise.

(e) Contiguous house price changes:-

$\Delta clrhp_r$ = Log change in real house prices regions contiguous to region r. The contiguity weights are based on the full time wage bills in contiguous regions.

(f) Other variables:-

cci = Index of credit conditions. Source: Fernandez-Corugedo and Muellbauer (2005).

lpc = Log consumer expenditure deflator. Source: ONS Blue Book.

rabmr = abmr - Δlpc = Real mortgage rate.

$pop2039_r$ = (population aged 20 to 39) / (population aged 20 to 69) in region r. Source: ONS.

ΔlrFTSE = change in log (FTSE/pc) i.e. real FTSE index. Source: ONS Financial Statistics.

ΔlrFTSEneg = ΔlrFTSE if negative and zero otherwise. It is a proxy for downside risk in the stock market.

sdr = average Stamp Duty rate. Source: HM Revenue & Customs online Stamp Duty statistics, Inland Revenue Statistics and own calculation).

D88 = dummy for mix of 1988 effects - replacement of domestic rates by Poll Tax and pre-announcement of the end of multiple mortgage interest tax relief on August 1st 1988.

D01 = dummy for 9/11 bombing and stock market turmoil in 2001.

(g) Data Descriptions and Unit Root Properties:-

We examined the unit root properties of the data and found few surprises. For example, real house prices and real income per house are both I(1); the tax adjusted mortgage rate abmr is also I(1); the change in the log of the real FTSE index is I(0) etc. Tables of descriptive statistics and various univariate and panel unit root and stationarity test results are available from the authors on request.