MODELLING STRUCTURAL CHANGE IN THE UK HOUSING MARKET:

A COMPARISON OF ALTERNATIVE HOUSE PRICE MODELS

BY

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Abstract

This paper develops a new approach to the modelling of house prices in the UK, with housing demand being conditioned directly on consumers' expenditure rather than the determinants of expenditure. Conditioning on consumption ensures that the permanent income measure used in determining the level of consumption is consistently reflected in housing demand. The effects of financial liberalisation on the relative consumption of housing and non-housing goods and services are captured using the average loan-value ratio for first-time buyers. We also allow for forward-looking behaviour via the expected real user cost of home ownership. House prices are assumed to adjust so as to clear the housing market. The proposed model is found to have structurally stable parameters across the housing market downturn since 1990. Statistical comparisons with the more conventional models in use at HM Treasury and the Bank of England during the early 1990s provide additional evidence in favour of our proposed approach.

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I. Introduction

Marked changes have occurred within the UK housing market over the last 15 years. In the early 1980s widespread financial deregulation raised the availability of mortgage finance and stimulated the demand for housing. Real house prices rose by over 4½ per cent per annum on average during the decade, with nominal house price inflation peaking at 28 per cent in 1988. Monetary policy was subsequently tightened in the late 1980s as inflationary pressures began to emerge in the wider economy, with mortgage interest rates rising to over 15 per cent in 1990. This helped to push the economy into recession, and the housing market weakened considerably. Nominal house prices fell by around 7½ per cent between 1990 and 1992, the first such decline for nearly forty years. The recent downturn is notable not only for the depth of the decline in real house prices, but also for its relative longevity, with real prices falling by an average 6¼ per cent per annum from 1989 to 1994 and by a further 2½ per cent in 1995. Movements in nominal and real house prices over the past three decades are shown in Chart 1.

Although these developments have led to considerable financial difficulties for many home-owners, with rising repossessions and widespread negative equity, they provide fertile ground for empirical research. In particular it is of interest to ask whether it is possible to obtain a model of house prices which has stable parameters across both the upturn of the 1980s, at a time of financial liberalisation, and the subsequent recession. Our findings suggest that stable relationships can be obtained, particularly if careful attention is paid to the appropriate specification of a structural housing demand relationship. However there is some evidence that prices were weaker in 1994 than might otherwise have been expected.

Many existing studies of house prices can be seen as being derived from the relationship between the desired quantity of housing services and the real user cost of those services (Poterba, 1984). In equilibrium the price paid by consumers for the flow of services from the housing stock, the real rental price of housing services, should equal the real user cost, which is the opportunity cost of investing in those services. Under the dual assumptions that both the quality and quantity of the housing stock are fixed in the short-term, real house prices must adjust to ensure equilibrium. The real rental price is unobservable, and is typically proxied using factors such as disposable incomes, demographic indicators and real interest rates which are viewed as determining the demand for housing services. Although the resulting model of house prices can then be given a structural interpretation, with the house price equation viewed as an inverted housing demand function, such models are also close to the reduced form specifications that could be derived from

equating particular housing demand and supply schedules and solving for house prices (Meen, 1990a).

The specification adopted in this paper differs from previous work in its treatment of housing demand. Whereas the conventional approach is to condition this on some measure of income, the specification developed here models the housing demand decision conditional on consumption itself. It is well known that an intratemporal condition equating the real user cost of housing to the marginal rate of substitution between the consumption of goods and the consumption of housing services can be derived from a standard utility maximising model of personal sector behaviour with two goods, housing services and a composite consumer good excluding housing (Westaway, 1992). We argue that this relationship can be used to obtain directly the long-run demand for housing, with house prices adjusting to remove disequilibrium between ex-ante demand and supply.

Conditioning on consumption rather than the determinants of consumption offers two important advantages. First, it ensures that the permanent income measure used in determining the level of consumption is consistently reflected in housing demand, so that consumption of goods and housing services cannot diverge indefinitely. This has obvious advantages in the context of wider macroeconometric models which typically contain existing relationships to determine consumers' expenditure. Second, it implies that the effects of financial liberalisation on the ratio of the consumption of non-housing goods and services to the consumption of housing services can be estimated separately from its common influence on their level.

We illustrate that the empirical results obtained using the theoretically correct measures of consumers' expenditure excluding housing services and the associated price deflator can be readily approximated by a model using total consumers' expenditure and the aggregate consumers' expenditure deflator. This simplifies the comparison of our suggested model of house prices with re-estimated versions of the models in use at the Bank of England (Breedon and Joyce, 1993) and the Treasury (HM Treasury, 1992) at the time of the housing market downturn. Both our model and that of the Bank have a common theoretical underpinning, treating the long-run house price equation as an inverted demand function. In contrast the Treasury model is best thought of as a reduced form equation, as it does not contain a housing stock measure, although it has some features in common with the other two models.

The rest of the paper is as follows. In Section II we derive our basic model of housing demand and discuss the treatment of financial liberalisation. Our econometric results using a data set constructed using data at 1985 prices are described in Section III. In

Section IV we re-estimate the two competing models and conduct a number of encompassing tests. In Section V we report preliminary results from re-estimating our preferred model on an updated and extended data set at 1990 prices. In Section VI we investigate the role of expectations in the determination of housing demand and house prices. Whilst there is general agreement that the real user cost of housing will depend on expectations of future house price movements, there is little agreement on the way in which expectations should be modelled. Concluding comments are given in Section VI, along with estimates of the direct impact of recent changes in the taxation of housing on house prices..

II A 'Conditional' Model Of Housing Demand.

The model of housing demand that we adopt in this paper arises out of the micro-to-macro model of household behaviour developed in Westaway (1992). In that model consumers derive utility from three different sources, the consumption of non-durable goods, the consumption of services from durable goods and the bequest that is left to successors. The consumer's lifetime utility function is assumed to be given by [1], for a (certain) lifespan of T years[1].

$$U(c,h,b) = \sum_{i=1}^{T}\left[\frac{U_i(c,h)}{(1+\rho)^i}\right] + \frac{U\left(b_{T+1}\right)}{(1+\rho)^{T+1}} \qquad [1]$$

where:
$$U_i(c,h) = \frac{c_i^{1-\gamma} + (\Theta h_i)^{1-\gamma}}{1-\gamma} \qquad [1a]$$

and:
$$U\left(b_{T+1}\right) = \frac{\alpha\,(W_{T+1} + g_{T+1}h_T)^{1-\gamma}}{1-\gamma} \qquad [1b]$$

Here c_i and h_i denote consumption excluding payment for actual and imputed housing services and the stock of housing assets, both defined in volume terms, held at age i, ρ is the rate of time preference, b_i denotes the bequest left to successors, W is total (non-housing) financial wealth, g denotes nominal house prices (PH) relative to consumer

[1] This embodies an implicit assumption of weak separability between the intertemporal allocation of consumption and the intratemporal split between goods and housing services.

prices (PCX)[2] and $1/\gamma$ is the intertemporal elasticity of substitution. The flow of housing services is assumed to be proportional to the stock held, with Θ denoting the constant of proportionality. Details of the equation that governs the evolution of wealth are given in Westaway (1992). Related models of household utility maximisation are considered by Dougherty and Van Order (1982), Meen (1990a) and Miles (1993).

The utility maximisation problem faced by the consumer involves the choice of c_i and h_i, for i=1..T, to maximise [1]. In the absence of any borrowing constraints the first order conditions from the resulting Lagrangean are:

$$\left[\frac{\delta L}{\delta c_i}\right] = \frac{(c_i)^{-\gamma}}{(1+\rho)^i} - \frac{\alpha'(1+r)^{T-1}}{(1+\rho)^T}\left[W_T + \frac{h_T\, g_{T+1}}{(1+r)}\right]^{-\gamma} \qquad [2a]$$

$$\left[\frac{\delta L}{\delta h_i}\right] = \frac{\Theta^{1-\gamma}(h_i)^{-\gamma}}{(1+\rho)^i} - \frac{\alpha' x_i(1+r)^{T-i}}{(1+\rho)^T}\left[W_T + \frac{h_T\, g_{T+1}}{(1+r)}\right]^{-\gamma} \qquad [2b]$$

where: $$\alpha' = \frac{\alpha(1+r)^{1-\gamma}}{1+\rho} \qquad [2c]$$

and $$x_i = \frac{g(r-\dot{g})}{1+r} \qquad [2d]$$

Here r denotes the real rate of interest, $(\dot{})$ denotes the time derivative of a variable and x_i represents the real user cost of housing. The user cost term reflects the difference between the present real cost of a unit of housing and the discounted expected cost one period hence.[3] Using the first-order conditions [2a] and [2b] it is possible to show that consumption and housing demand are linked by the intratemporal condition [3], with the marginal rate of substitution between the two determined by the real user cost of housing.

$$\frac{h_i}{c_i} = (x_i)^{-\frac{1}{\gamma}}\Theta^{\frac{1-\gamma}{\gamma}} \qquad [3]$$

Aggregating across consumers and rearranging yields an expression for house prices of the form:

[2] PCX is the implicit price deflator for the consumption of non-housing goods and services.

[3] We make the simplifying assumption that the depreciation of housing assets and other property taxes are zero. This is relaxed in the empirical work below.

$$g = \frac{PH}{PCX} = \frac{(CX/H)^{\gamma}\,\Theta^{1-\gamma}}{(r-\dot{g})/(1+r)} \qquad [4]$$

where CX and H denote total consumption excluding housing services and the aggregate stock of (owner-occupied) dwellings.[4]

If binding borrowing constraints are present, the effective user cost of housing will be increased, distorting the standard relationship between the consumption of goods and housing services. However Westaway (1992) illustrates that the impact of changes in borrowing constraints on [3] is considerably smaller than the impact on the overall level of either consumption expenditure or housing demand, largely because the 'income' effect of financial liberalisation affects both forms of expenditure.[5] This suggests that a possible empirical model of house prices can be developed from [4], possibly modified by some allowance for borrowing conditions, since conditioning on consumers' expenditure provides an explicit link between the permanent income measure implicit in both consumption and housing demand.

In contrast, many previous studies of house prices have started from [4], or an equivalent capital market condition using the real rental return on housing, but substituted in the determinants of consumption, so that prices depend upon measures of income and wealth.[6] As Meen (1990a) suggests, the resulting house price equations are very close to the reduced form specifications that could be obtained by substituting in the separate determinants of housing demand and supply. Such equations may also modify the link between the permanent income measure implicit in the consumption function and housing demand.

Dougherty and Van Order (1982) argue that [3] cannot be used to derive estimates of housing demand since consumption is endogenous. This is certainly true if the objective is to derive ex-ante forecasts of demand or undertake policy analysis, although the same could be said of models conditioned on income, particularly if the number of hours worked is a choice variable. For estimation purposes, a key issue is whether consumption can be regarded as weakly exogenous; if it is, then efficient estimation and testing can be conducted by analysing the conditional model (for housing demand or house prices) and

[4] One possible extension of this model would be to allow for secondary properties owned by the personal sector that are either part of the rental market or unoccupied.

[5] The simultaneous regime shifts in consumers' expenditure and housing demand as a result of financial liberalisation provide an example of 'co-breaking' (Hendry, 1996).

[6] Recent examples include Dicks (1990), Meen (1990b) and Breedon and Joyce (1993).

neglecting the marginal model (for consumption and the real interest rate). Ericsson (1992) and Urbain (1992) suggest that one means of testing for weak exogeniety in a systems context is to estimate any cointegrating vectors that may exist and test for their significance in the marginal model. If the vector is insignificant, the respective variable can be regarded as weakly exogenous within the conditional model. A test of whether consumption can be regarded as weakly exogenous in [4] is reported below. If consumption is endogenous then it is still possible to estimate an equation of the form we propose, but only within a wider system jointly modelling consumers' expenditure and housing demand.

In practice some allowance still has to be made for the unobservable shadow price associated with any binding borrowing constraint. Dougherty and Van Order (1982) suggest that the effect of credit market constraints can differ according to whether there are absolute constraints on the volume of loans or implicit constraints on the ratio of loans to collateral, ('disequilibrium' and 'equilibrium' credit rationing in the terminology of Kent, 1980). Both forms of constraint have been used by financial institutions in the UK. Quantity constraints were clearly present in the mortgage market in the 1970s, with building societies rationing mortgage supply either through queuing or by giving priority to existing depositors. The cartel operated through the Building Societies Association at that time kept average mortgage rates below other money market rates. Increased competition following financial deregulation and the entry of the banks into the mortgage market in 1980 resulted in borrowing rates moving closer to market clearing levels, although demand could still be affected through prudential changes in non-price criteria for loans, such as the implicit loan-value ratio offered to new borrowers. As Chart 2 illustrates, the loan-value ratio has continued to fluctuate since financial deregulation and was tightened over the course of the recent recession.

A number of different measures of mortgage rationing have been used in recent studies of the housing market, all of which attempt to capture the overall impact of rationing rather than the impact on housing demand relative to consumption. Hendry (1984) uses the outstanding stock of mortgage debt while Meen (1990a,b) employs an explicit estimate of the extent of mortgage rationing. Dicks (1990) uses a measure of mortgage rationing constructed by Wilcox (1985) from the loan-value ratio for first-time buyers. Muellbauer and Murphy (1992) use both the lagged mortgage stock and a financial liberalisation index derived from the behaviour of the loan-value ratio.

A priori it is far from clear how to best capture changes in the mortgage market. In the empirical work below, we investigate three separate proxies. The first, also used in HM Treasury (1992), is the difference between the three-month interbank rate and the average

mortgage rate. We expect this variable to have a negative effect, with positive values arising at times of mortgage rationing. There was little margin between the two interest rates in the 1980s, with wholesale markets often being the marginal source of funds for many building societies.

The other two measures make use of the loan-value ratio for first time buyers.[7] Conventional tests suggest that this series has a unit root, reflecting the marked upward jump in the early 1980s following the onset of financial deregulation. However allowing for the presence of a one-time shift in the mean of the series from the fourth quarter of 1981, using the model of Perron (1990), serves to reject the unit root hypothesis in favour of a deterministic shift in the series at this point.[8] The average loan-value ratio is estimated to have risen from 78.3 per cent in the period of rationing to 84 per cent in the deregulated market. The post-rationing average loan-value ratio was taken as an estimate of the level at which excess demand was zero. (A related approach is adopted by Wilcox, 1985.) Two rationing proxy variables, denoted LVR1 and LVR2, were then constructed, with LVR1=84-LVR and LVR2 = 78.3+(5.7*DU)-LVR, where LVR denotes the actual loan-value ratio and DU is a dummy equal to 1 from 1982Q1 onwards. Again, both variables are expected to have a negative impact on housing demand.[9]

All three of the financial liberalisation measures we investigate are stationary variables. This implicitly assumes that financial deregulation may have done little to disturb permanently the long-run relationship between the growth of housing demand and the growth of consumers' expenditure. This finding is consistent with the predictions of the micro-to-macro work and the suggestion that this effect of financial liberalisation was likely to be less important than those which affected the level of both categories of demand.

In contrast, it is likely that changes in the desire for owner-occupation arising from either wider social and demographic trends or from the 'right-to-buy' legislation introduced in

[7] The loan-value ratio for first time buyers provides a better indicator of lending criteria than the overall loan-value ratio. The latter also reflects changes in the desire of existing owner-occupiers to hold capital gains in the form of housing equity (Wilcox, 1985).

[8] A similar finding is reported in Breedon and Joyce (1993), although they date the break in the fourth quarter of 1980.

[9] The interest differential and the loan-value measures have a sample correlation coefficient of 0.6 over the period from 1967 to 1981. However the correlation between them and the measure of mortgage rationing in Meen (1990a) is negligible.

the 1980s,[10] will raise the demand for housing services relative to the growth of consumers' expenditure over time. Changes in the age distribution of the population may also affect the demand for housing to the extent that this varies between different age groups. Poterba (1991) provides a useful discussion of the role of demographic factors. A number of UK studies have tested for the effects of such changes by including age-specific terms in house price models. Dicks (1990), Milne (1991) and Breedon and Joyce (1993) all use the proportion of the population aged 25-29, on the grounds that the demand for housing services typically rises sharply within this age group. Dicks (1990) and Meen (1990a) also include (different) measures for the number of households, although this can be viewed as a necessary counterpart to the use of per capita income and the total stock of dwellings.

In principle such an adjustment should not be required in our model. This is because [3] implies a long-run coefficient of unity on consumption in the housing demand relationship; thus it does not matter whether the model is specified in per capita terms or not, since the population terms will cancel. However, as our quantity indicator is the stock of owner-occupied dwellings, we may need to allow for any factors which serve to raise the stock of such dwellings relative to consumption. We thus include the owner-occupation rate as an additional regressor.[11]

A modified version of equation [4] forms the starting point for the long-run housing demand equation, or equivalently the long run house price equation that we estimate. Taking logarithms and re-arranging yields:

$$\ln(g) = f\,(\ln(C), \ln(H), U, Z) \qquad [5]$$

Here U denotes the user cost of housing and Z denotes the additional variables included to take account of financial liberalisation and changes in the rate of owner-occupancy. The user cost term modifies the real interest rate term in [4] to allow for the tax deductibility of mortgage interest payments, property taxes and housing depreciation:

$$U = \frac{USER - (\dot{P}H)^e}{1 + .01 * \left[USER - (\dot{P}C)^e\right]} \qquad [6]$$

[10] The 'right-to-buy' legislation allowed many existing council house tenants to purchase their houses at a discounted price. Arguably some of these tenants would otherwise have remained within the rented sector.

[11] If we assume that the consumption of owner-occupiers (C_o/O) is proportional to aggregate consumption per capita (C/P), with λ denoting the ratio of the consumption of an individual house-

where: $USER = \alpha(1-t_m)RM + (1-\alpha)(1-t_i)RB + \kappa + \delta + \tau$ [6a]

Here α denotes the average loan-value ratio, t_m is the effective rate of tax relief on mortgage interest payments, t_i is the standard rate of income tax, RM and RB denote mortgage and base interest rates, κ denotes the property tax rate, δ is depreciation and τ is transactions costs. The property tax rate was derived using the ratio of household sector rates (including water rates) to the value of the housing stock. Detailed information on δ and τ is harder to obtain; we have assumed that they average 0.9 percentage points per annum. There are a number of ways in which price expectations can be handled. For now, we simply use the actual change in house prices and consumer prices over the past year.[12] In practice we find that [6] is occasionally negative, in contrast to what might be expected given [3], and thus the user cost cannot be entered in logarithmic form in [5].

One possible extension to [6] would be to abandon the implicit certainty equivalence assumption and allow for the possibility that consumers may build a risk premium into the user cost expression to account for the risky nature of house price returns. Miles (1993) suggests that a risk premium of 2-8 per cent (per annum) be added to the user cost, based on the average, actual excess return on UK equities over a 'safe' rate of interest over the period from 1919-1988. However whilst this ensures that annual measures of the user cost are non-negative, it does not perform the same service for the quarterly measure, and we do not pursue the issue here.

Disequilibrium in the housing market is given by the imbalance between the existing stock of housing assets and the long-run 'desired' demand [ln(H)*] which can be obtained by inverting [5]. Given the assumption that house prices act in a market clearing fashion, one possible approach is to simply include estimates of short-term 'housing disequilibrium' (HDIS), derived from [5], as an error-correction term in a dynamic house price equation of the form:

$$\Delta\ln(g) = A(L)\Delta\ln(g)_{-1} + \lambda(HDIS_{-1}) \qquad [7]$$

where L denotes the lag operator (Δ=1-L) and HDIS = ln(H) - ln(H)*.

The long-run housing demand equation can be estimated using cointegration techniques. Use of Johansen's (1988) multivariate procedure allows a direct test of whether there is an

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owner to the average, then: $\ln(H/O)-\ln(C_o/O) = \ln(H/O)-\ln(\lambda.C/P) = \ln(H)-\ln(C)-\ln(\lambda)-\ln(O/P)$. This implies that the owner-occupation rate should have a long-run elasticity of unity.

[12] The results of Muth (1986) provide some evidence in favour of an adaptive expectations approach. We explore this issue in more detail in Section VI below.

unique cointegrating vector between the variables in [5]. If not, it is then difficult to interpret [5] as an inverted long-run housing demand equation.

Alternatively, it is possible that a more efficient estimate of the parameters in the long-run housing demand curve can be obtained by estimating the long-run structure jointly with the dynamics of the house price equation. Kremers *et al*.(1992) argue that such an approach can have greater power in situations where there is an unique cointegrating vector. However, if this approach is to be used to recover a long-run demand equation, it does require that the specification search is not extended to include either additional non-stationary variables or variables that affect the supply-side of the housing market.

III. Aggregate house prices - The NIESR Model

In this section we report the empirical results obtained using a model in which house prices, and hence housing demand, are conditioned on consumers' expenditure. Two sets of results are reported. The first use a measure of consumers' expenditure that excludes housing services and the corresponding implicit consumer price deflator. The second investigates whether similar results can in fact be obtained using total consumers' expenditure and the aggregate consumers' expenditure deflator. Both approaches are found to yield similar results, suggesting that any bias arising from the use of total expenditure is not especially significant. In both cases we begin by reporting the results obtained using the Johansen procedure, before exploring the findings from more conventional single equation approaches. We estimate three separate models for each measure of consumers' expenditure in an attempt to discriminate between the three suggested proxies for financial liberalisation.[13] All models are initially estimated on a sample from 1968Q1-1990Q4, with nine additional observations retained for a post-sample stability test.[14] All variables are defined in the Data Appendix.

The starting point of the Johansen procedure is to specify an unrestricted vector autoregression (VAR) that provides a statistically adequate description of the data. Three endogenous variables were used, the ratio of housing assets to non-housing consumers' expenditure (H/CX), the owner-occupation rate (OO) and real house prices (PH/PCX).

[13] When all three measures were included in the same model they were jointly significant, but individually insignificant. Use of a single indicator per model enables a more informative comparison to be made between the indicators as well as between the other models we re-estimate.

[14] As we use a data set initially constructed by Pain and Westaway (1994) using UK National Accounts data at 1985 prices we do not have additional data for all the variables used beyond 1993Q1.

All are non-stationary on the basis of conventional unit root tests. Three separate models are reported; all include a measure of the real user cost (RUSERX) plus one of the financial liberalisation indicators as additional exogenous regressors.[15] Initial inspection of the equations in the unrestricted VAR led to the inclusion of a number of dummy variables in order to ensure that the resulting residuals were normally distributed. The chosen order of the VAR was the minimum required to 'pass' conventional serial correlation tests. The resulting diagnostic statistics for a fifth-order VAR are reported in Table 1.

Four dummy variables are included in each model. Three, denoted D68, DUM73 and DUM79, are for outliers arising from the impact of changes in indirect taxation in 1968, 1973 and 1979 on the pattern of consumption (see Pain and Westaway, 1994). The third measure, denoted DUM, is a composite variable based on the three separate dummies used in our earlier work. It takes the value 1 in 1972Q3 and 1988Q3 and -1 in 1971Q2. The 1988 dummy is probably picking up speculative effects in the housing market resulting from the impact of the pre-announced imposition of restrictions on the value of mortgage interest tax relief in the March 1988 Budget.[16] The other dummies may possibly reflect temporary changes in demand pressures arising as a result of changes in controls on bank lending in 1971 and 1972.

There appears to be relatively little difference between the respective models. In all cases the Johansen cointegration test statistics suggest the presence of a single cointegrating vector. Normalising this vector on the ratio of housing demand to consumption gives an equation in which relative house prices have a negative elasticity close to -0.5 and the owner-occupancy term has an elasticity a little above unity. The imposition of a unit elasticity on the latter is readily accepted by the data.[17]

If the reported cointegrating vector is a housing demand curve, and house prices adjust to remove disequilibrium in the housing market, it should be possible to enter the residuals from the cointegrating vector as an error-correction term in a house price equation. For all

[15] The real user cost of housing is stationary over our sample period [ADF(2)=-4.19].

[16] An alternative might be to try and capture this using the cubic 'frenzy' terms in house prices employed by Hendry (1984) and others, although we are sceptical that these types of effect add much more than a simple dummy variable.

[17] All three models in Table 1 use the ratio of the housing stock to consumers' expenditure. The parameter restriction implied by this term is accepted by the data. For example, if Model 1 is re-estimated with separate housing and consumption terms, the cointegration tests both point to a single cointegrating vector, and the restriction required to give the ratio of the two terms is readily accepted [Chi(1)=2.26]. The residuals from this restricted vector are not significant in the consumers expenditure equation in the VAR [Chi(1)=2.96].

three models the residuals from the cointegrating vector were significant in the house price equation in the VAR and insignificant in the other two equations. For Model 2 the respective test statistics of discarding the residuals term were [Chi(1)=22.5] in the house price equation and [Chi(1)=0.23] and [Chi(1)=0.95] in the other two equations. This implies that the owner-occupation rate and the housing stock/consumption ratio may be regarded as weakly exogenous to real house prices.

As the Johansen procedure suggests that a single cointegrating vector is present, it is possible to estimate a dynamic house price equation directly. This has the additional advantage that it obviates the need to hunt for dummy variables so as to obtain satisfactory reduced form equations for consumption and the owner-occupation rate in the VAR. We began with an unrestricted fifth order distributed lag model with allowance being made for possible separate dynamic effects from the nominal user cost and the housing stock. The initial step was to test whether the dynamic terms in the housing stock and the owner-occupation rate could be dropped, since it was felt that any reported effects could be spurious given the need for interpolation to obtain quarterly data. (This is an additional reason for not using the Johansen approach alone.) For all three models these dynamic terms were found to be jointly insignificant.

The final data-based restricted dynamic equations over the sample to 1990Q4 are reported in the first three columns of Table 2. In all cases the elasticities on the housing stock term and both consumers' expenditure and owner-occupancy are equal and opposite to one another. All three restrictions are readily accepted by the data. The real user cost term is well determined and there are significant effects from the separate financial liberalisation terms. The implied long-run parameters at the foot of the table indicate that a one per cent rise in consumers expenditure may ultimately raise real house prices by some 2-2¼ per cent, whilst a one percentage point increase in the real user cost may reduce prices by some 1¼-1¾ per cent.

There is relatively little to choose between the three alternative models. Conventional encompassing tests were unable to discriminate between the models, suggesting that all three financial liberalisation measures are capturing similar effects. All have similar dynamics and satisfy conventional diagnostic tests, including a test for parameter stability over 1991 and 1992. Potentially this is a powerful finding, given that the behaviour of nominal house prices over these two years was quite different to that at any time within our sample period. There was little difference in any of the coefficients when the equations were re-estimated up until 1993Q1. Column 4 in Table 2 reports the parameters obtained from the re-estimated version of (2.2), using the LVR1 variable.

The inverted long-run solutions to the house price equations are also shown at the foot of Table 2. The long-run price elasticities are close to those obtained in Table 1 using the Johansen approach and the price elasticity of -0.5 reported in Meen (1990a). These estimates can be used to calculate the implied extent of 'disequilibrium' between housing supply and demand. In (2.4) for example, 'disequilbrium' is given by:

$$HDIS = \ln(H) - \ln(CX) - \ln(OO) + 0.4833*\ln(PH/PCX) + 0.00882*RUSERX \quad [8]$$

The resulting series, normalised to average zero, is shown in Chart 3, with 'desired demand' exceeding supply in the early and late 1970s and the second half of the 1980s. It is interesting to note that the most recent downturn appears to have been accompanied by an unusually prolonged period of 'excess supply', especially as compared to the downturn in the early 1980s. However the extent of disequilbrium across the most recent cycle is estimated to have been a little weaker than across the cycle in the early 1970s.

In Tables 3 and 4 we report the results of re-estimating the models reported in Tables 1 and 2 with alternative measures of consumers' expenditure. Whilst it is theoretically appropriate to use a measure of consumers' expenditure that excludes expenditure on housing services, this is not especially convenient for many macroeconometric models. Typically, such models either contain equations for expenditure on durable goods and non-durable goods and services (including housing services), or a single equation for total consumers' expenditure. Consumer prices are usually modelled with a single equation for the overall consumers' expenditure deflator. The alternative models we wish to compare against our suggested approach both use the overall consumers' expenditure deflator to construct real house prices. We thus investigate whether the use of aggregate consumers' expenditure imparts any significant bias into our model. This entails using two new variables, total consumption (C) and the aggregate consumer price deflator (PC), with a corresponding re-definition of real house prices and the real user cost of housing.[18]

Overall, the results obtained are close to those reported in Tables 1 and 2. A unit elasticity on owner-occupation can again be imposed in the Johansen estimates, with an implied price elasticity of -0.45 to -0.5 per cent, although the unrestricted coefficients are now smaller than in the original model. The housing stock/ consumption ratio again appeared weakly exogenous with respect to real house prices. For Model 2 the respective test

[18] It could be argued that an IV estimator be employed to overcome any simultaneity arising from the use of total consumers' expenditure. However this will not necessarily deal with the measurement error in the dependent variable (real house prices) resulting from the use of the aggregate consumers' expenditure deflator.

statistics of discarding the cointegrating vector residuals were [Chi(1)=26.4] in the house price equation and [Chi(1)=0.70] and [Chi(1)=0.66] in the other two equations.

The dynamic equations reported in Table 4 have similar properties to those in Table 2, although their performance over the post-sample period appears a little worse. The implied effects of consumers' expenditure and the real user cost of house price are in line with those in Table 2, and the long-run parameters from Table 2 can be readily imposed if desired. For example, the test statistics of imposing the long-run parameters from (2.2) and (2.4) in (4.2) and (4.4) are [Chi(2)=0.65] and [Chi(2)=0.53] respectively. One implication of this finding is that the long-run properties of the consumption of housing services are similar to the long-run properties of the consumption of other goods and services, so that few differences result from using a measure including both categories of expenditure.

IV. A Comparison With Alternative House Price Equations

IV.1 Structural Models

As the approach we adopt to modelling house prices is somewhat unconventional, it is of interest to compare our results with those that can be obtained from the alternative specifications recently proposed by Breedon and Joyce (1993) and HM Treasury (1992). In both cases we obtain data-consistent versions of their equations on a sample ending in 1990Q4, and test for post-sample parameter stability.

Breedon and Joyce (1993) use the Johansen technique to estimate a long-run cointegrating vector for house prices, with prices related to real personal disposable income (Y), the real gross financial assets of the personal sector (RGW), the stock of owner-occupied dwellings and the proportion of the adult population aged 25-29 (P25TO29). Additional terms in real interest rates (RUSER1), net liquid assets (RNL), the loan-value ratio for first time buyers and changes in repossessions (REP) are contained in their dynamic price equation. Further details are provided in the Data Appendix. In using this framework we have made one change, using a one quarter backward-looking measure of inflation in the real user cost term rather than a one period forward-looking measure of inflation expectations.[19]

[19] This change can be justified as our main interest lies in the comparison of the statistical properties of the competing models rather than their responses in policy simulations. Expectational effects are likely to be subsumed within the data-based dynamics of the re-estimated equation. We return to this

In practice it did not prove possible to estimate a VAR with Gaussian errors with the explanatory variables used in the Bank specification, even after a hunt for suitable dummies. (The main problem lies in finding an adequate reduced form equation for the interpolated population share variable.) Results from a fourth-order VAR in which the real user cost is included as an additional exogenous regressor are reported in Table 5. The income and housing stock elasticities are 1.93 and -0.89, compared to 2.87 and -2.13 in Breedon and Joyce (1993). The restrictions required to obtain their cointegrating vector were rejected using a fourth-order VAR [Chi(4)=23.61)] and a second-order VAR (as used by Breedon and Joyce). However, since there is some evidence that there may be an unique long-run relationship, we proceeded to estimate a unrestricted dynamic house price equation.

The final restricted dynamic equation is reported in the first column of Table 6.[20] The long-run solution is broadly consistent with the findings in Table 5, although the relative size of the real wealth elasticity has increased. A data-acceptable unit long-run elasticity has been imposed on the housing stock. The real user cost term again appears to be important, although the long-run parameter is smaller than those reported in Tables 2 and 4. There are additional significant dynamic effects from changes in the level of repossessions (measured relative to the housing stock), the nominal user cost, and the growth of real liquid assets and real disposable income. However the loan-value indicator (LVR2) was not found to be significant. Although this model includes a much wider range of variables than those in Table 4, the overall fit is only marginally better within sample. There is also clear evidence of parameter instability over the post-sample period.

The results from re-estimating (6.1) over the full sample period are shown in (6.2). The standard error has risen by over 10 per cent and a number of terms have become insignificant. One possible factor behind this is the behaviour of the chosen demographic variable; whilst the movement in the cohort aged 25-29 happens to accord with the observed changes in house prices over the late 1980s, it does not appear to have done so in the early 1990s.

IV.2 Reduced Form Models

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issue below, and explicitly test for the presence of forward-looking expectations in the user cost term of our preferred equation.

[20] The sample period is different to that used in Tables 2 and 4. This is because the absence of data on repossessions meant that the general model could only be estimated from 1969Q2; however it would be possible to estimate the final model over a longer sample period.

The HM Treasury (1992) house price equation is best regarded as a reduced form model, even though it contains many terms that would be present if it were purely an inverted demand function. This is because the model does not contain any quantity indicator, as might be expected of a structural demand relationship. It is not therefore possible to recover the parameters of the housing demand schedule. Prices are determined by real disposable non-property income per capita, the corresponding net financial wealth to income ratio, real interest rates defined using consumer price inflation (REAL) and the spread between the three month interbank rate and the mortgage rate which we have already used (RS-RM). The latter variable is a proxy for financial liberalisation effects. The real interest rate differs a little from that used in our model, as it uses both an alternative cost of housing finance variable and consumer, rather than house price, inflation. The interest rate on housing finance differs from [6] in that it includes an estimate of the opportunity cost of funds invested in home ownership, but does not allow for depreciation and property taxes. In practice the two nominal user cost measures are closely correlated.[21]

The results here use aggregate real disposable income rather than non-property income, simply because the latter has to be constructed using a number of proxies for the labour share of the income of the self-employed. Comparison between the results reported here and those in Pain and Westaway (1994) suggests that similar results are obtained using either measure of income.

Initial attempts at re-estimating the dynamic equation in H M Treasury (1992) over the original sample period (1969Q1-1990Q2) proved unsuccessful. Whilst the majority of variables remained significant, the standard error of the re-estimated equation was 2.2 per cent compared to 1.57 per cent in the original, and there was evidence of serial correlation. However, use of the Johansen procedure did suggest that real house prices, income per head (Y/POP) and the wealth-income ratio (RNW/Y) generated a single cointegrating vector, see Table 7. Real interest rates and the financial liberalisation proxy were included in the VAR as additional stationary regressors, along with a number of additional dummies to allow for outliers in the wealth-income ratio. The resulting income elasticity of 1.49 is a little below that of 1.7 in HM Treasury (1992).

Again, we proceeded to estimate a data-based dynamic model of house prices. The preferred equation is reported as (6.3) in Table 6, and is estimated up until 1990Q4. The fit of the equation is comparable to the other models, and in contrast to the Bank model

[21] The sample correlation coefficient over 1968Q1-1993Q1 is 0.963. We are grateful to Rod Whittaker for supplying us with the Treasury measure.

the post-sample stability test is passed. The implied long-run income elasticity is 1.37 per cent, and the real interest rate effect is larger than in any of the other reported models over this sample period. The estimates imply that a 1 percentage point rise in real interest rates will reduce real house prices by 1¾ per cent. The financial liberalisation proxy also appears to be better determined than in (4.1). The results from re-estimating (6.3) up until 1993 are reported in (6.4); there is a small decline in the income elasticity and a marginal deterioration in the fit of the equation.

IV. 3 Encompassing Tests

For each model it appears possible to find a well determined house price equation across the estimation sample, although it is clear that the alternative model based on the Bank of England specification exhibits post-sample structural instability. Table 8 reports the results from a number of encompassing tests of the respective competing models.[22] The first panel reports the variable deletion tests required to return to each model from the general 'artificial regression' (MacKinnon, 1992) formed by combining all of the separate regressors in each model. Over the shorter sample period each model is accepted against the general model. However over the extended sample the restrictions required to obtain the Bank model are rejected, implying that the other two models contain information that would improve its' specification.

Equivalent pairwise tests are reported in the other two panels of Table 8, using the JA-test of Fisher and McAleer (1981) and variable deletion tests from an artificial regression model. Overall, the findings provide some evidence that our model conditioned on consumption is to be preferred to the other competing models, since it encompasses both models over at least one sample period but is not encompassed by either at conventional significance levels, over either sample period.

V. Results With An Updated Data Set

All our empirical results so far have been obtained from a data set containing National Accounts data at 1985 prices, and hence we have not been able to explore whether any of the reported house price equations can successfully capture the longevity of the recent downturn as well as its depth. In this section we report further results from an updated data set using the latest National Accounts data at 1990 prices. We are able to report

[22] The test statistics were calculated using equations estimated from 1969Q2, in line with the sample period for the re-estimated Bank model.

estimates up to the end of 1994, reflecting the last published data points for both owner-occupation and the owner-occupied housing stock.

The initial results from re-estimating our preferred model on the new data set are reported in Table 9. In contrast to the data set at 1985 prices there was clear evidence of seasonal effects in the 1990 based data set, and hence we include seasonal dummies in all the reported equations.[23]

In the first column we re-estimate equation (4.2), augmented by seasonal dummies, over the same sample from 1968Q1 to 1990Q4. The resulting parameter estimates are similar to those of (4.2), with the exception of the lagged dynamic term in real house prices. Discarding this term and the insignificant seasonal dummy for the first quarter of the year gives equation (9.2). The implied long-run relationships are little changed from those of (4.2), with consumers' expenditure having an elasticity of 2 per cent and the real user costs having a semi-elasticity of 1.4 per cent. The rebasing of consumers' expenditure to 1990 prices appears to have made some difference to the dynamics, with a much bigger coefficient being obtained on the contemporaneous term. The post-sample Chow test again fails to reject the null of parameter stability, although it was apparent that the forecasting 'performance' over the sample including 1994 is not as good as over the previous three years of the downturn.

This can be seen more clearly from equation (9.3) where we re-estimate the model using data up until the end of 1994. All the parameters remain significant, but there is some evidence of a decline in the overall goodness of fit, and the last five residuals from the equation were all negative, reflecting a tendency to over-predict the growth of real house prices. There is also some evidence that a number of individual parameters have changed significantly. For example the parameters on the user cost term and the loan-value ratio have both become larger and better determined as the sample is extended, with the semi-elasticity on the user cost rising to 2 per cent. This reflects the fact that these two terms are having to account for the delayed recovery in house prices, with the user cost being raised by restrictions to the value of mortgage interest tax relief and the loan-value ratio reflecting the tighter prudential criteria being applied by many lenders. In contrast, the growth of both consumers' expenditure and the owner-occupation rate across 1993 and 1994 was acting to raise long-term housing demand. Overall, these preliminary results

[23] There is some weak evidence of seasonal effects in the 1985 based data, although it is not significant. For example seasonal dummies were jointly insignificant when added to both (2.2) and (4.2) with respective test statistics [F(3,80)=1.94] and [F(3,80)=1.66]. However they were jointly significant when (4.2) was re-estimated on the 1990 based data [F(3,80)=7.96].

suggest that it would be worthwhile to undertake further research into the factors that lie behind the unusual longevity of the recent downturn in the housing market.

Finally, it is of interest to test whether our financial liberalisation measures can fully account for the extent of mortgage rationing prior to the 1980s. We have investigated two alternative, widely used, proxies. The first is the mortgage rationing measure of Meen (1990a,b), which we set to zero after 1981 following HM Treasury (1992). The second is the real value of the outstanding stock of mortgages, as used in Hendry (1984), Dicks (1990) and Muellbauer and Murphy (1992). The latter two studies continue to include this variable even though their samples extend into the 1980s when quantity rationing was minimal. To test the significance of these terms, (9.3) was augmented with first to fourth order lagged terms in the variables. In both cases the resulting parameters were jointly insignificant, with respective test statistics of [F(4,94)=0.42] for mortgage rationing and [F(4,94)=1.49] for the real mortgage stock. A separate measure equal to the mortgage stock until 1981 and zero thereafter also proved to be jointly significant, albeit marginally so [F(4,94)=2.42].

However one of the lagged mortgage stock terms was significant, as shown in equation (9.4). This additional variable appears to be orthogonal to the other variables in the model, with the possible exception of the loan-value ratio, which is found to have a larger effect than before. In fact the actual long-run effect from the mortgage stock on house prices is small, and insignificant at conventional levels, with an elasticity of only 0.0238 per cent and a standard error of 0.0125 per cent. The existence of this term implies that whilst the loan-value ratio and consumption capture most of the financial constraints on housing demand prior to deregulation, there may be a specific element of quantity rationing which is not fully reflected in either. This suggests that it could be of interest to investigate further the precise timing and extent of deregulation.

VI. The Role Of Expectations

Whilst there is general agreement that the real user cost of house purchase will depend on expectations of future house price movements, there is little agreement on the way in which such expectations should be modelled. Actual expectations of future house prices are not available and thus some proxy measure has to be used. The simplest procedure is to substitute current or past inflation as a proxy for future expectations in the user cost [6]. Alternatively it is possible to use predicted values from a regression of current house price inflation on a more extensive lagged information set. Hendry (1984) adopts this approach and argues that agents should be assumed to have 'sensible' expectations that are not fully efficient rather than rational expectations, since the latter impose excessive informational requirements. Meen (1990a) reports that both of the approaches in which expectations are proxied by a measure based on past behaviour yield similar results. More generally, the ex-post future rate of house price inflation can be used in place of the ex-ante expectation. This gives rise to an errors in variables problem and so an instrumental variables technique has to be utilised.

In order to investigate the means by which house price expectations are formed we respecify the real user cost term [6] to allow for a weighted average of forward and adaptive expectations of the rate of change of house prices. We assign a weight β to the future annual rate of house price inflation and a weight $(1-\beta)$ to the current annual rate of inflation, with the resulting user cost term denoted RUSERF. (The measure of the user cost used thus far is one with β set to zero.) Further details are provided in the Data Appendix.

Substitution of the ex-post outturn for the ex-ante expectation of house price inflation means that an instrumental variables technique needs to be employed. We initially employ a standard, non-linear instrumental variables estimator, and subsequently re-estimate the preferred equation using a variant of the Hayashi-Sims estimator allowing for the potential existence of a third-order moving average error term.[24]

With the user cost term being lagged one period, house price expectations are implicitly being formed at time t-1. Hence current dated instruments are not appropriate, since they may be correlated with the forecasting error (the difference between actual and expected prices). We therefore exclude the contemporaneous dynamic term in consumers' expenditure from the instrument set. A number of additional instruments were used; these

[24] In general k-step ahead rational expectations forecast errors are serially correlated with a moving average process of order (k-1). Here we substitute the actual outturn in place of the ex-ante expected price four quarters ahead, thus generating a potential third-order MA error.

consisted of lagged values of the exogenous variables plus an extrapolative predictor of house price inflation derived from a time series equation using the chain rule of forecasting.[25] Although the generated regressor could be substituted in place of the ex-ante expectation, this would not yield consistent estimates of the variance-covariance matrix.

The resulting equations are reported in Table 10. Full details of the instruments used are given at the foot of the Table. The first equation is estimated over the shorter sample period to the end of 1990, whilst the second extends the sample to 1994. In both cases the fit of the equation appears a little better than the corresponding estimates in Table 9 with $\beta=0$ (equations (9.2) and (9.3) respectively). Current house price inflation receives a weight of approximately three-quarters. The forward-looking measure has a weight of approximately one-quarter, and is significantly different from zero over both sample periods. However it is not particularly well determined and a wide range of values could be imposed. For example the hypothesis of equal weights on future and current inflation ($\beta=0.5$) is not rejected in (10.2), [Wald(1)=2.93].

The long-run parameters at the foot of the table indicate that the explicit allowance for forward expectations has raised the estimated long-run impact of the real user cost on housing demand and house prices by nearly fifty per cent, although the long-run semi-elasticity is not as well determined as before. Thus in equation (10.2) for example, a sustained one percentage point rise in the real user cost is estimated to reduce real house prices by 3 per cent. The remaining parameters in the model appear to be little changed by the allowance for forward-looking expectations, with the elasticity of house prices with respect to consumers' expenditure remaining around 2 per cent.

These estimates can again be used to calculate the implied extent of 'disequilibrium' between housing supply and demand. In (10.2) for example, 'disequilbrium' is given by:

$$HDIS = \ln(H) - \ln(C) - \ln(OO) + 0.5186*\ln(PH/PC) + 0.0158*RUSERF \qquad [9]$$

The resulting series, normalised to average zero, is shown in Chart 4. This has a similar pattern to that shown in Chart 3. It can be seen that the model suggests that the prolonged

[25] The equation was: $\Delta\ln(PH) = \underset{(5.7)}{0.486}\Delta\ln(PH)_{-1} + \underset{(4.6)}{0.410}\Delta\ln(PH)_{-2} - \underset{(4.2)}{0.314}\Delta\ln(PH)_{-5}$ + seasonals

Sample Period:1968Q1-90Q4; standard error 1.80%; LM(4)=1.20; CHOW(16)=14.43.

It is of some interest to note that this simple time series equation also appears to exhibit parameter stability between 1990 and 1994. Letting the basic regression be $x_t = \Sigma\alpha_j(x_{t-j})$, the chain rule of forecasting implies that the current expected value of x_{t+i} is given by ${}_tx^e_{t+i} = \Sigma_{j=1}\ \alpha_j({}_tx^e_{t+i-j})$

downturn in the housing market has been accompanied by a prolonged period of 'excess supply'. However more recent data for 1995 suggest that 'excess demand' may have begun to appear once more[26], although the implied extent of disequilbrium remains below that seen in the late 1980s.

One potential problem with equations (10.1) and (10.2) is that they do not allow for the possible presence of a third-order moving average error, although the LM test statistics suggest there is little evidence of any significant serial correlation. Equation (10.3) allows for the presence of a moving average error using an IV technique in Pesaran and Pesaran (1991) closely related to the Hayashi-Sims technique. Given the null of serial correlation we only include fourth and fifth order lags of variables in the instrument set. This restricts the available sample size to one beginning in 1968Q3. The results provide some support for the validity of the previous regressions without an explicit moving average error, as none of the error terms is individually significant and they are jointly insignificant [Wald(3)=3.04]. The implied long-run parameters are little changed from (10.2).

VII. Conclusions

In this paper we have presented a new approach to the determination of house prices, with housing demand being conditioned directly on consumers' expenditure rather than the determinants of expenditure. This ensures that the first-order condition which relates the marginal rate of substitution between the consumption of housing services and the consumption goods to the real user cost of housing will be satisfied. This theoretical consistency is particularly convenient in a wider macromodelling context, since most macroeconometric models will already have an existing relationship that determines consumers' expenditure. Indeed the approach we propose has recently been introduced into both the National Institute Domestic Econometric Model and the new HM Treasury Model (Chan *et al*., 1995), although there are some important differences in that the NIESR model equation has forward-looking expectations in the real user cost term, whereas the Treasury version is entirely backward-looking..

An explicit allowance has been made for the presence of forward-looking behaviour via price expectations in the econometric work, in contrast to many previous studies of house prices in the UK. We have also sought to explore whether a variety of different

[26] Data for some of the variables used in our model are not available for 1995 at the time of writing, and so values from the May 1996 NIESR forecast base have been used in the calculation of HDIS after the end of 1994.

approaches to modelling house prices can explain both the depth and the longevity of the recent downturn in the housing market. In general it appears that the scale of the decline in house prices in 1991 and 1992 can be better explained than the subsequent stagnation of nominal prices across 1993 and 1994. Comparisons with other models provide some evidence in favour of our new model.

The results obtained here can be used to estimate the direct impact of recent changes in the provision of mortgage interest tax relief on house prices. Such changes will affect the real user cost of housing. They will of course also affect consumers' expenditure, as disposable income will change and because house price movements affect the value of the net wealth of the personal sector (Pain and Westaway, 1994). These effects are likely to enhance the direct effects working through the user cost.

At the end of 1994 we estimate that the real user cost was some 1.1 percentage points higher than it would have been had the effective rate of mortgage interest tax relief remained at the 1987 level, all other things being equal. Since that time the effective rate of tax relief is estimated to have declined from 27 per cent to 11¼ per cent. This reflects three factors; a failure to uprate the threshold below which interest payments are eligible for tax relief, the gradual reduction in the rate at which tax relief can be claimed and the limitation of tax relief to one claimant per household. The long-run semi-elasticity of -3 percentage points in equation (10.2) implies that the change in mortgage interest tax relief may have reduced real house prices by 3¼ per cent, for given expenditure, since 1987.[27]

However some of this may have been offset by other changes in the structure of the public finances since the late 1980s. In particular the move from domestic rates to the community charge and then the council tax, has also affected the user cost. Domestic rates were paid by all owner-occupiers, but not by tenants unless the rates paid by private landlords were capitalised in rental charges. In contrast the community charge was effective on all adults and the council tax is normally based on residency, so that it is paid by the owner of an owner-occupied property and the tenant of a rented property. Thus the abolition of domestic rates is likely to have reduced the user cost of owner-occupation, since owner-occupiers would remain liable to tax in the rented sector. Our data suggests that changes in the effective rate of property charges have reduced the real user cost by 0.9 percentage points since 1987, offsetting most of the direct effects of the reduced rate of mortgage interest tax relief. However changes in property charges will not have offset any of the additional income-related effects of restrictions in mortgage interest relief.

[27] At the start of 1995, real house prices were approximately 4 per cent below their average level in 1987.

The proposed model has deliberately been kept as simple as possible in order to highlight its salient features. There is some weak evidence that further account could be taken of the extent of quantity restrictions on mortgage lending in the 1970s. Equally it might be possible to investigate the role of more recent developments such as the rise in repossessions and mortgage arrears, and the number of borrowers with negative equity. This will require further work on the theoretical foundations of the housing market; at present such factors are not directly incorporated in the typical utility maximising model at the heart of most housing demand models.

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Chart 1: Annual House Price inflation (per cent)

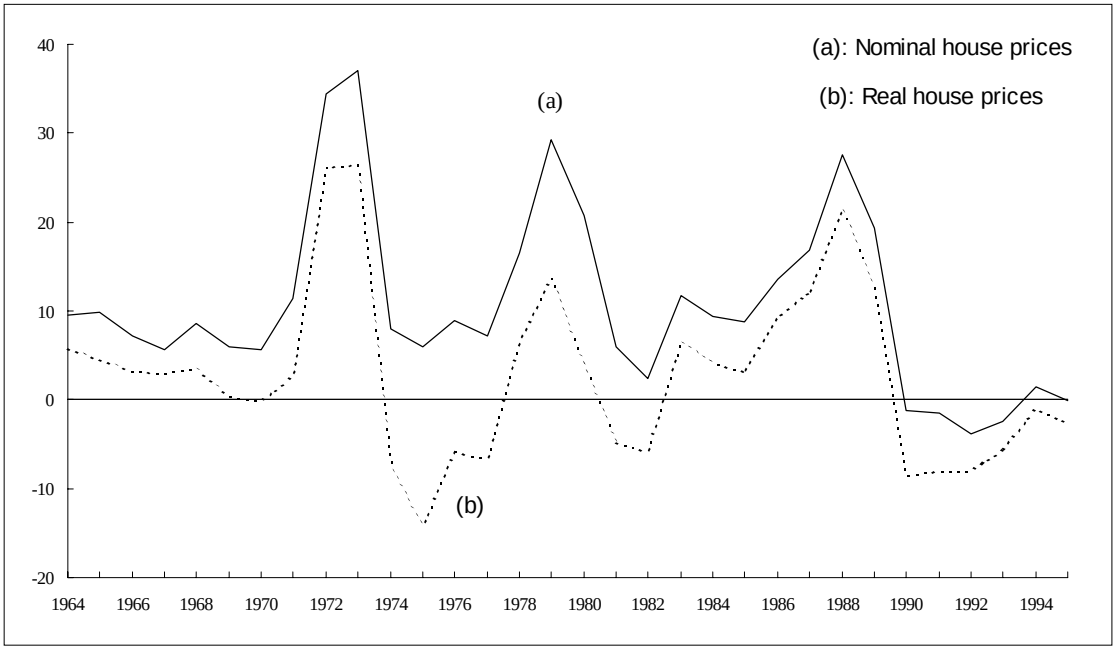
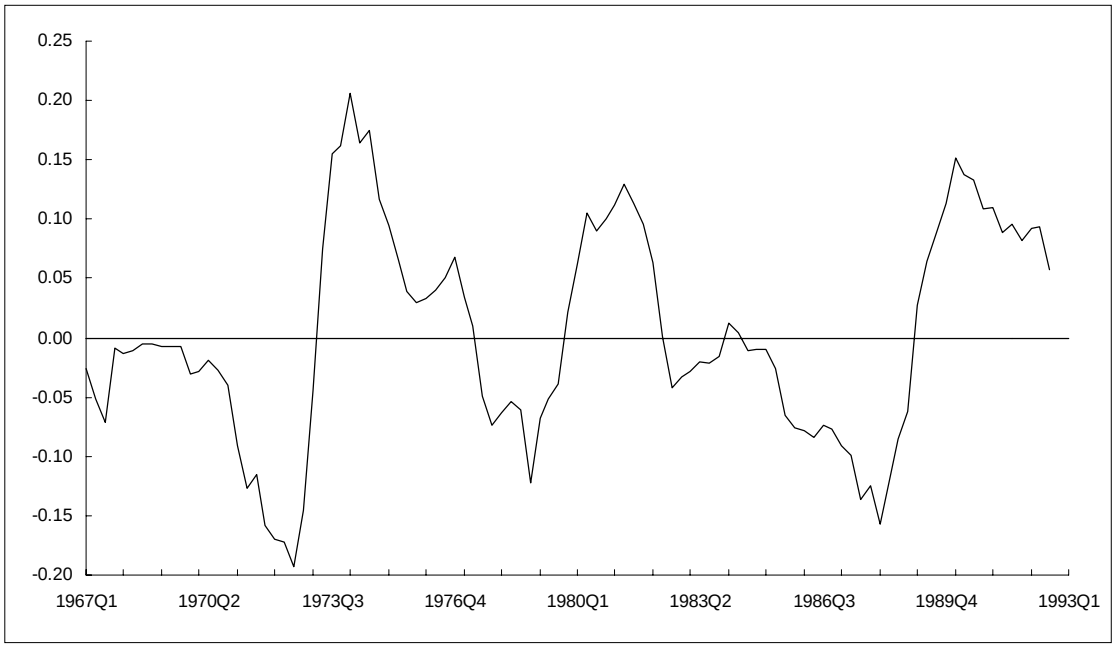


Chart 2: Average Loan-Value Ratio, First-Time Buyers (per cent)

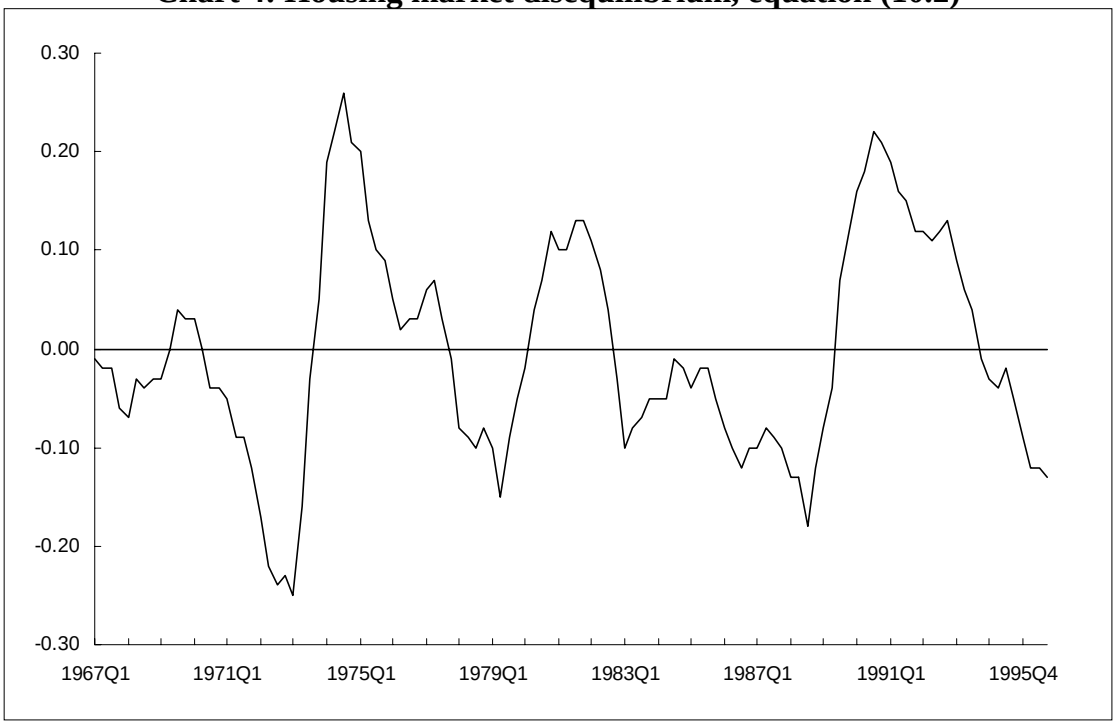


Chart 3: Housing market disequilibrium, equation (2.4)



Note: a negative figure indicates 'excess' demand

Chart 4: Housing market disequilibrium, equation (10.2)



Note: a negative figure indicates 'excess' demand

Table 1: Johansen Estimates of Housing Demand, NIESR Model

Sample: 1968Q1-1990Q4; VAR(5) Model

Exogenous regressors: Model 1: $RUSERX_{-1}$, $(RS-RM)_{-1}$, DUM, D682, D73, D79
Model 2: $RUSERX_{-1}$, $LVR1_{-1}$, DUM, D682, D73, D79
Model 3: $RUSERX_{-1}$, $LVR2_{-1}$, DUM, D682, D73, D79

Diagnostic Statistics for the VAR						
Variable	**Serial Correlation**	**Normality**	**Heteroscedasticity**			
Model 1						
ln(H/CXH)	F(4,66)=1.46	Chi(2)=0.28	F(1,90)=0.38			
ln(OO)	F(4,66)=2.23	Chi(2)=6.15	F(1,90)=0.01			
ln(PH/PCX)	F(4,66)=1.72	Chi(2)=1.46	F(1,90)=0.30			
Model 2						
ln(H/CXH)	F(4,66)=1.71	Chi(2)=0.34	F(1,90)=0.39			
ln(OO)	F(4,66)=0.97	Chi(2)=8.63	F(1,90)=0.01			
ln(PH/PCX)	F(4,66)=0.66	Chi(2)=1.40	F(1,90)=0.86			
Model 3						
ln(H/CX)	F(4,66)=1.68	Chi(2)=0.28	F(1,90)=0.66			
ln(OO)	F(4,66)=1.27	Chi(2)=9.84	F(1,90)=0.05			
ln(PH/PCX)	F(4,66)=0.63	Chi(2)=1.08	F(1,90)=0.28			
Cointegration test based on maximal eigenvalue of the stochastic matrix						
Null	Alternative	Test Statistic			95% Critical Value	
		Model 1	Model 2	Model 3		
r = 0	r = 1	22.63	26.15	24.26	20.97	
r = 1	r = 2	4.55	6.88	4.91	14.07	
Cointegration test based on trace of the stochastic matrix						
Null	Alternative	Test Statistic			95% Critical Value	
		Model 1	Model 2	Model 3		
r = 0	r = 1	27.48	33.05	31.90	29.68	
r = 1	r = 2	4.81	6.90	7.63	15.41	
Cointegrating Vectors (normalised on housing demand)						
Variable	Estimated			Restricted		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
ln(H/CX)	-1.0	-1.0	-1.0	-1.0	-1.0	-1.0
ln(OO)	1.1522	1.1194	1.3976	1.0	1.0	1.0
ln(PH/PCX)	-0.5382	-0.5216	-0.5666	-0.4895	-0.4845	-0.4435
Test of restrictions	**Model 1**	Chi(1)=0.28				
	Model 2	Chi(1)=0.16				
	Model 3	Chi(1)=2.35				

Table 2. Dynamic Equations for NIESR Models

Dependent Variable: Δln(PH/PCX)

	(2.1)	**(2.2)**	**(2.3)**	**(2.4)**
Constant	-1.9479 (6.0)	-1.6848 (4.3)	-1.9292 (5.2)	-1.7061 (4.6)
$\ln(PH/PCX)_{-1}$	-0.1696 (6.3)	-0.1555 (5.3)	-0.1569 (4.9)	-0.1563 (5.5)
$\ln(H/CX)_{-1}$	-0.3673 (6.0)	-0.3194 (4.3)	-0.3642 (5.2)	-0.3233 (4.7)
$\ln(OO)_{-1}$	0.3673 (6.0)	0.3194 (4.3)	0.3642 (5.2)	0.3233 (4.7)
$RUSERX_{-1}$	-0.0023 (9.9)	-0.0029 (8.3)	-0.0027 (7.9)	-0.0029 (8.9)
$\Delta\ln(PH/PCX)_{-2}$	-0.1305 (2.4)	-0.2265 (2.4)	-0.2044 (2.1)	-0.2046 (2.3)
$\Delta\ln(PH/PCX)_{-4}$	0.1305 (2.4)			
$\Delta\ln(CX)$	0.4051 (3.3)	0.3822 (3.1)	0.4229 (3.5)	0.3636 (3.1)
$\Delta USER_{-1}$	-0.0095 (2.2)	-0.0109 (2.7)	-0.0121 (3.0)	-0.0120 (3.2)
DUM	0.0620 (5.8)	0.0640 (6.0)	0.0627 (5.8)	0.0638 (6.1)
$(RS-RM)_{-1}$	-0.0039 (2.5)			
$LVR1_{-1}$		-0.0022 (2.8)		-0.0021 (2.8)
$LVR2_{-2}$			-0.0026 (2.2)	
Sample	1968Q1-90Q4	1968Q1-90Q4	1968Q1-90Q4	1968Q1-93Q1
Diagnostic statistics				
$\overline{R^2}$	0.7378	0.7373	0.7291	0.7403
SE	0.0179	0.0179	0.0181	0.0176
LM [Chi(4)]	2.52	2.59	1.55	2.92
BJ [Chi(2)]	7.66	2.46	2.67	2.28
ARCH [Chi(4)]	3.08	4.77	2.73	4.46
CHOW [Chi(9)]	5.20	6.22	5.91	
Test of restrictions on general model				
	F(23,60)=0.82	F(23,60)=0.79	F(23,60)=0.90	
Long-run parameters (house prices)				
ln[(CX*OO)/H]	2.1657 (22.3)	2.0543 (14.5)	2.3210 (19.8)	2.0691 (17.6)
RUSERX	-0.0133 (6.1)	-0.0187 (4.8)	-0.0172 (4.7)	-0.0182 (5.3)
Long-run parameters (housing demand)				
ln(CX*OO)	1.0 (-)	1.0 (-)	1.0 (-)	1.0 (-)
RUSERX	-0.0061 (5.8)	-0.0091 (4.0)	-0.0074 (5.1)	-0.0088 (4.6)
ln(PH/PCX)	-0.4618 (22.3)	-0.4868 (14.5)	-0.4309 (19.8)	-0.4833 (17.6)

Notes: t-statistics in parentheses.

Table 3: Johansen Estimates, NIESR Model using aggregate consumption

Sample: 1968Q1-1990Q4; VAR(5) Model

Additional exogenous regressors: Model 1: $RUSER_{-1}$, $(RS-RM)_{-1}$, DUM, D682, D73, D79
Model 2: $RUSER_{-1}$, $LVR1_{-1}$, DUM, D682, D73, D79
Model 3: $RUSER_{-1}$, $LVR2_{-1}$, DUM, D682, D73, D79

Diagnostic Statistics for the VAR						
Variable	**Serial Correlation**	**Normality**	**Heteroscedasticity**			
Model 1						
ln(H/C)	F(4,66)=1.52	Chi(2)=0.21	F(1,90)=0.40			
ln(OO)	F(4,66)=2.50	Chi(2)=6.43	F(1,90)=0.02			
ln(PH/PC)	F(4,66)=1.98	Chi(2)=1.50	F(1,90)=0.27			
Model 2						
ln(H/C)	F(4,66)=1.76	Chi(2)=0.51	F(1,90)=0.39			
ln(OO)	F(4,66)=1.10	Chi(2)=9.06	F(1,90)=0.03			
ln(PH/PC)	F(4,66)=0.96	Chi(2)=1.08	F(1,90)=0.23			
Model 3						
ln(H/C)	F(4,66)=1.68	Chi(2)=0.46	F(1,90)=0.67			
ln(OO)	F(4,66)=1.42	Chi(2)=10.5	F(1,90)=0.08			
ln(PH/PC)	F(4,66)=0.92	Chi(2)=0.86	F(1,90)=0.28			
Cointegration test based on maximal eigenvalue of the stochastic matrix						
Null	Alternative	Test Statistic			95% Critical Value	
		Model 1	Model 2	Model 3		
r = 0	r = 1	26.95	32.09	28.90	20.97	
r = 1	r = 2	3.06	6.26	2.15	14.07	
Cointegration test based on trace of the stochastic matrix						
Null	Alternative	Test Statistic			95% Critical Value	
		Model 1	Model 2	Model 3		
r = 0	r = 1	30.09	38.34	32.05	29.68	
r = 1	r = 2	3.14	6.26	3.15	15.41	
Cointegrating Vectors (normalised on housing demand)						
Variable	Estimated			Restricted		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
ln(H/C)	-1.0	-1.0	-1.0	-1.0	-1.0	-1.0
ln(OO)	0.8948	0.8506	1.1140	1.0	1.0	1.0
ln(PH/PC)	-0.4560	-0.4387	-0.4932	-0.4967	-0.4939	-0.4517
Test of restrictions	**Model 1**	Chi(1)=0.25				
	Model 2	Chi(1)=0.50				
	Model 3	Chi(1)=0.36				

Table 4. Dynamic Equations for NIESR Models with aggregate consumption

Dependent Variable: $\Delta\ln(PH/PC)$

	(4.1)	**(4.2)**	**(4.3)**	**(4.4)**
Constant	-2.4146 (7.0)	-2.1659 (4.8)	-2.4413 (6.3)	-2.0478 (4.9)
$\ln(PH/PC)_{-1}$	-0.2081 (7.4)	-0.1930 (6.0)	-0.1988 (6.1)	-0.1847 (6.0)
$\ln[H/(C*OO)]_{-1}$	-0.4389 (6.9)	-0.3950 (4.9)	-0.4442 (6.3)	-0.3736 (5.0)
$RUSER_{-1}$	-0.0025 (11.2)	-0.0032 (9.4)	-0.0031 (9.0)	-0.0032 (9.8)
$\Delta\ln(PH/PC)_{-2}$	-0.1486 (2.8)	-0.2565 (2.8)	-0.2390 (2.6)	-0.2326 (2.6)
$\Delta\ln(PH/PC)_{-4}$	0.1486 (2.8)			
$\Delta\ln(C)$	0.3974 (2.8)	0.3774 (2.7)	0.4096 (2.9)	0.3400 (2.4)
$\Delta USER_{-1}$	-0.0091 (2.3)	-0.0099 (2.6)	-0.0108 (2.8)	-0.0108 (3.0)
DUM	0.0618 (6.1)	0.0634 (6.3)	0.0623 (6.1)	0.0635 (6.3)
$(RS-RM)_{-1}$	-0.0029 (2.0)			
$LVR1_{-1}$		-0.0018 (2.2)		-0.0018 (2.4)
$LVR2_{-2}$			-0.0021 (1.9)	
Sample	1968Q1-90Q4	1968Q1-90Q4	1968Q1-90Q4	1968Q1-93Q1
Diagnostic statistics				
$\overline{R^2}$	0.7570	0.7546	0.7510	0.7498
SE	0.0170	0.0170	0.0172	0.0171
LM [Chi(4)]	5.16	5.26	4.10	4.86
BJ [Chi(2)]	8.95	4.39	4.56	2.52
ARCH [Chi(4)]	3.88	4.63	3.13	3.78
CHOW [Chi(9)]	11.68	9.81	12.54	
Test of restrictions on general model				
	F(23,60)=0.66	F(23,60)=0.72	F(23,60)=0.80	
Long-run parameters (house prices)				
ln[(C*OO)/H]	2.1095 (25.9)	2.0474 (17.1)	2.2279 (24.5)	2.0222 (18.7)
RUSER	-0.0122 (7.8)	-0.0166 (5.8)	-0.0149 (6.3)	-0.0171 (6.2)
Long-run parameters (housing demand)				
ln(C*OO)	1.0 (-)	1.0 (-)	1.0 (-)	1.0 (-)
RUSER	-0.0058 (7.3)	-0.0081 (4.8)	-0.0069 (6.6)	-0.0085 (5.3)
ln(PH/PC)	-0.4741 (25.9)	-0.4884 (17.1)	-0.4476 (24.4)	-0.4945 (18.7)

Notes: t-statistics in parentheses.

Table 5: Johansen Estimates, Bank of England Model

Sample: 1969Q2-1990Q4; VAR(4) Model

Additional exogenous regressors: Model 1: $RUSER1_{-1}$, $LVR2_{-1}$, DUM

Diagnostic Statistics for the VAR			
Variable	**Serial Correlation**	**Normality**	**Heteroscedasticity**
ln(Y)	F(4,59)=2.28	Chi(2)=2.60	F(1,85)=2.04
ln(RGW)	F(4,59)=2.47	Chi(2)=9.06	F(1,85)=0.15
ln(PH/PC)	F(4,59)=2.21	Chi(2)=0.29	F(1,85)=1.43
ln(H)	F(4,59)=1.56	Chi(2)=5.07	F(1,85)=0.29
P2529	F(4,59)=2.06	Chi(2)=275.3	F(1,85)=0.05
Cointegration test based on maximal eigenvalue of the stochastic matrix			
Null	Alternative	Test Statistic	95% Critical Value
r = 0	r = 1	31.01	33.46
r = 1	r = 2	23.31	27.07
Cointegration test based on trace of the stochastic matrix			
Null	Alternative	Test Statistic	95% Critical Value
r = 0	r = 1	74.54	68.52
r = 1	r = 2	43.54	47.21
Cointegrating Vectors (normalised on real house prices)			
ln(PH/PC) = 1.9333 ln(Y) - 0.8893 ln(H) + 0.0154 ln(RGW) + 0.1080 P2529			

Table 6. Dynamic Equations for Bank and Treasury Models

Dependent Variable: $\Delta\ln(PH/PC)$

	(6.1)	**(6.2)**	**(6.3)**	**(6.4)**
Constant	-3.1873 (7.2)	-2.2510 (4.6)	-0.0376 (3.5)	-0.0318 (3.0)
$\ln(PH/PC)_{-1}$	-0.2364 (7.9)	-0.1814 (5.3)	-0.1353 (5.9)	-0.1207 (5.4)
$\ln(H)_{-1}$	-0.2364 (7.9)	-0.1814 (5.3)		
$\ln(Y)_{-1}$	0.3908 (6.7)	0.2816 (4.1)		
$\ln(Y/POP)_{-1}$			0.1866 (5.0)	0.1551 (4.4)
$\ln(RGW)_{-1}$	0.0750 (4.9)	0.0645 (4.1)		
$(RNW/Y)_{-1}$			0.0056 (3.2)	0.0045 (2.7)
$P2529_{-1}$	0.0150 (2.7)	0.0023 (0.3)		
$RUSER1_{-1}$	-0.0019 (3.2)	-0.0023 (3.4)		
$REAL_{-1}$			-0.0024 (4.4)	-0.0021 (4.0)
$\Delta\ln(PH/PC)_{-1}$	-0.4261 (1.9)	-0.5047 (1.9)	0.3716 (5.7)	0.3695 (5.7)
$\Delta\ln(PH/PC)_{-3}$	0.2109 (3.7)	0.2326 (3.2)	0.2095 (3.0)	0.2169 (3.1)
$\Delta\ln(PH/PC)_{-4}$	0.2733 (3.5)	0.2535 (3.1)	0.2911 (3.9)	0.3020 (4.1)
$\Delta\ln(Y)$	0.2595 (2.8)	0.2413 (2.4)		
$\Delta\ln(Y/POP)$			0.1995 (1.9)	0.1912 (1.8)
$\Delta USER_{-1}$	-0.0107 (3.5)	-0.0114 (3.5)		
ΔREP	-0.6576 (2.3)	-0.1884 (0.6)		
$\Delta\ln(RNL)_{-2}$	0.2085 (2.1)	0.1843 (1.6)		
DUM	0.0650 (17.8)	0.0730 (16.7)	0.0675 (6.5)	0.0677 (6.4)
$(RS-RM)_{-1}$			-0.0059 (4.2)	-0.0063 (4.5)
Sample Period	1969Q2-90Q4	1969Q2-93Q1	1968Q1-90Q4	1968Q1-93Q1
Diagnostic statistics				
$\overline{R^2}$	0.7776	0.7278	0.7562	0.7456
SE	0.0166	0.0183	0.0170	0.0173
LM [Chi(4)]	7.37	9.93	6.53	6.10
BJ [Chi(2)]	1.43	5.02	0.30	2.07
ARCH [Chi(4)]	8.38	2.57	5.25	2.42
CHOW [Chi(9)]	26.0		11.91	
Test of restrictions on general model				
	F(24,49)=0.69		F(20,61)=0.64	
Long-run parameters				
ln(H)	-1.0 (-)	-1.0 (-)		
ln(Y)	1.6533 (17.2)	1.5526 (10.9)		
ln(Y/POP)			1.3789 (13.9)	1.2849 (12.4)
ln(RGW)	0.3175 (4.48)	0.3558 (3.43)		
RNW/Y			0.0412 (3.31)	0.0370 (2.78)
P2529	0.0633 (2.67)	0.0125 (0.35)		
RUSER1	-0.0082 (2.84)	-0.0129 (2.99)		
REAL			-0.0174 (4.75)	-0.0173 (4.20)

Notes: t-statistics in parentheses.

Table 7: Johansen Estimates, HM Treasury Model

Sample: 1968Q1-1990Q4; VAR(5) Model

Additional exogenous regressors: Model 1: $REAL_{-1}$, $(RS-RM)_{-1}$, DUM, D874, DUMW

Diagnostic Statistics for the VAR			
Variable	**Serial Correlation**	**Normality**	**Heteroscedasticity**
ln(Y/POP)	F(4,66)=0.64	Chi(2)=2.34	F(1,90)=0.10
RNW/Y	F(4,66)=0.80	Chi(2)=0.41	F(1,90)=1.68
ln(PH/PC)	F(4,66)=1.02	Chi(2)=4.74	F(1,90)=0.01
Cointegration test based on maximal eigenvalue of the stochastic matrix			
Null	Alternative	Test Statistic	95% Critical Value
r = 0	r = 1	29.99	20.97
r = 1	r = 2	6.85	14.07
Cointegration test based on trace of the stochastic matrix			
Null	Alternative	Test Statistic	95% Critical Value
r = 0	r = 1	36.85	29.68
r = 1	r = 2	6.85	15.41
Cointegrating Vectors (normalised on real house prices)			
ln(PH/PC) = 1.4876 ln(Y/POP) + 0.0318 (RNW/Y)			

Table 8. Encompassing Test Statistics

Model 1: Equation 2, Table 4 Model 2: Equation 1, Table 6

Model 3: Equation 3, Table 6

Variable deletion test on general model			
Sample	**Maintained Model**		**Test statistic**
1969Q2-1990Q4	Model 1		F(16,62)=1.36 *[.19]*
	Model 2		F(11,62)=1.08 *[.40]*
	Model 3		F(14,62)=1.53 *[.13]*
1969Q2-1993Q1	Model 1		F(16,71)=1.25 *[.26]*
	Model 2		F(11,71)=2.31* *[.02]*
	Model 3		F(14,71)=1.50 *[.13]*
JA-test statistics			
Sample	**Maintained Model**	**Alternative Model**	**Test statistic**
1969Q2-1990Q4	Model 1	Model 2	JA = 1.32 *[.19]*
	Model 2	Model 1	JA = 1.77 *[.08]*
	Model 1	Model 3	JA = 1.99 *[.06]*
	Model 3	Model 1	JA = 2.27* *[.02]*
	Model 2	Model 3	JA = 1.82 *[.07]*
	Model 3	Model 2	JA = 1.68 *[.09]*
1969Q2-1993Q1	Model 1	Model 2	JA = 0.27 *[.79]*
	Model 2	Model 1	JA = 4.30 *[.00]*
	Model 1	Model 3	JA = 1.77 *[.08]*
	Model 3	Model 1	JA = 2.66* *[.01]*
	Model 2	Model 3	JA = 3.52 *[.00]*
	Model 3	Model 2	JA = -.20 *[.84]*
Pairwise variable deletion tests			
Sample	**Maintained Model**	**Alternative Model**	**Test statistic**
1969Q2-1990Q4	Model 1	Model 2	F(11,67)=1.40 *[.20]*
	Model 2	Model 1	F(6,67)=0.94 *[.47]*
	Model 1	Model 3	F(8,70)=1.40 *[.21]*
	Model 3	Model 1	F(6,70)=1.39 *[.23]*
	Model 2	Model 3	F(6,67)=1.91 *[.09]*
	Model 3	Model 2	F(9,67)=2.09* *[.04]*
1969Q2-1993Q1	Model 1	Model 2	F(11,76)=1.37 *[.21]*
	Model 2	Model 1	F(6,76)=3.43* *[.01]*
	Model 1	Model 3	F(8,79)=1.31 *[.25]*
	Model 3	Model 1	F(6,79)=1.60 *[.16]*
	Model 2	Model 3	F(6,76)=3.23* *[.01]*
	Model 3	Model 2	F(9,76)=1.44 *[.18]*

Notes: p-values in parentheses. A test statistic significant at the 5 per cent level indicates that the maintained model is rejected against the alternative model.

Table 9. Dynamic Equations for NIESR Models using an updated data set

Dependent Variable: $\Delta\ln(PH/PC)$

	(9.1)	**(9.2)**	**(9.3)**	**(9.4)**
Constant	-2.0685 (6.0)	-2.0706 (6.0)	-1.2290 (4.7)	-1.3848 (5.3)
$\ln(PH/PC)_{-1}$	-0.1723 (7.0)	-0.1724 (7.1)	-0.1170 (5.8)	-0.1104 (5.7)
$\ln[H/(C*OO)]_{-1}$	-0.3465 (5.9)	-0.3468 (6.0)	-0.2048 (4.7)	-0.2315 (5.2)
$RUSER_{-1}$	-0.0023 (7.8)	-0.0024 (11.4)	-0.0023 (12.0)	-0.0022 (12.5)
$\Delta\ln(PH/PC)_{-2}$	0.0285 (0.3)			
$\Delta\ln(C)$	0.5855 (4.5)	0.5760 (4.5)	0.5199 (4.3)	0.5412 (5.0)
$\Delta USER_{-1}$	-0.0083 (3.1)	-0.0085 (3.6)	-0.0088 (3.8)	-0.0093 (3.2)
DUM	0.0477 (5.0)	0.0479 (5.1)	0.0472 (4.6)	0.0449 (5.2)
$LVR1_{-1}$	-0.0012 (2.0)	-0.0013 (2.5)	-0.0023 (5.1)	-0.0034 (4.6)
Q1	-0.0024 (0.6)			
Q2	0.0130 (2.8)	0.0136 (3.6)	0.0131 (3.7)	0.0127 (3.8)
Q3	0.0187 (4.2)	0.0193 (5.7)	0.0203 (6.2)	0.0203 (6.0)
$\ln(SM81/PC)_{-2}$				0.0026 (2.2)
Sample	1968Q1-90Q4	1968Q1-90Q4	1968Q1-94Q4	1968Q1-94Q4
Diagnostic statistics				
$\overline{R^2}$	0.8343	0.8377	0.8151	0.8218
SE	0.0137	0.0136	0.0141	0.0139
LM [Chi(4)]	2.64	1.59	2.48	1.25
BJ [Chi(2)]	3.83	4.34	3.38	1.58
ARCH [Chi(4)]	0.36	0.40	0.60	1.08
Chow [Chi(16)]	22.63	23.59		
Long-run parameters (house prices)				
$\ln[(C*OO)/H]$		2.0118 (18.7)	1.7514 (13.9)	2.0978 (10.3)
RUSER		-0.0139 (6.8)	-0.0199 (6.5)	-0.0202 (6.3)
$\ln(SM81/PC)_{-2}$				0.0238 (1.9)
Long-run parameters (housing demand)				
$\ln(C*OO)$	1.0 (-)	1.0 (-)	1.0 (-)	1.0 (-)
RUSER		-0.0069 (5.5)	-0.0113 (5.2)	-0.0096 (5.4)
$\ln(PH/PC)$		-0.4971 (18.7)	-0.5708 (13.9)	-0.4767 (10.3)

Notes: t-statistics in parentheses.

Table 10. Expectations and House Prices

Dependent Variable: $\Delta\ln(PH/PC)$

	(10.1)	**(10.2)**	**(10.3)**
Constant	-1.5250 (3.2)	-0.9382 (3.4)	-1.1377 (4.5)
$\ln(PH/PC)_{-1}$	-0.1234 (3.4)	-0.0812 (3.3)	-0.0963 (4.5)
$\ln[H/(C*OO)]_{-1}$	-0.2555 (3.2)	-0.1567 (3.4)	-0.1905 (4.5)
$RUSERF_{-1}^{\dagger}$	-0.0026 (13.2)	-0.0025 (12.5)	-0.0027 (13.2)
β	0.7481 (5.9)	0.7161 (5.7)	0.7757 (9.0)
$(1-\beta)$	0.2519 (2.0)	0.2839 (2.3)	0.2243 (3.4)
$\Delta\ln(C)^{\dagger}$	0.5443 (2.7)	0.6064 (2.6)	0.3627 (3.0)
$\Delta USER_{-1}$	-0.0062 (2.0)	-0.0065 (2.2)	-0.0094 (3.7)
DUM	0.0435 (5.4)	0.0426 (5.1)	0.0456 (5.6)
$LVR1_{-1}$	-0.0014 (2.2)	-0.0019 (3.5)	-0.0019 (3.8)
Q2	0.0145 (4.4)	0.0139 (4.4)	0.0121 (4.3)
Q3	0.0208 (6.0)	0.0215 (6.6)	0.0205 (7.1)
Moving Average Error Parameters			
ε_{t-1}			-0.0623 (0.6)
ε_{t-2}			0.1314 (1.3)
ε_{t-3}			0.0828 (0.8)
Sample	1968Q1-1990Q4	1968Q1-1994Q4	1968Q3-1994Q4
Diagnostic statistics			
$\overline{R^2}$	0.8608	0.8433	0.8307
SE	0.0126	0.0130	0.0136
LM [Chi(4)]	3.63	2.58	n/a
BJ [Chi(2)]	4.34	3.77	n/a
SARGAN [Chi(8)]	7.37	10.10	n/a
Long-run parameters (house prices)			
ln[(C*OO)/H]	2.0710 (14.4)	1.9284 (9.9)	1.9791 (12.4)
RUSER	-0.0208 (3.1)	-0.0303 (3.1)	-0.0280 (4.2)
Long-run parameters (housing demand)			
ln(C*OO)	1.0 (-)	1.0 (-)	1.0 (-)
RUSER	-0.0100 (3.0)	-0.0158 (3.3)	-0.0142 (4.4)
ln(PH/PC)	-0.4829 (14.4)	-0.5186 (9.9)	-0.5053 (12.4)

Notes: t-statistics in parentheses. All equations estimated by instrumental variables. An † denotes a variable that has been instrumented in equations (10.1) and (10.2). Additional instruments were $\Delta_4\ln(PH)_{-1}$, $\Delta_4\ln(PH)_{-5}$, $\Delta\ln(C)_{-1}$, $\Delta\ln(C)_{-2}$, $\Delta\ln(PH/PC)_{-1}$, $\Delta(USER)_{-2}$, $LVR1_{-2}$, Q1, $USER_{-1}$ and a generated regressor derived from a time series equation for house prices, see text for details.

Equation (10.3) allows for a third-order moving average error. The instrument set used was Q_1, Q_2, Q_3, DUM, $\Delta_4\ln(PH)_{-5}$, $\Delta\ln(C)_{-4}$, $\Delta\ln(C)_{-5}$, $\Delta\ln(PH/PC)_{-4}$, $\Delta\ln(PH/PC)_{-5}$, $\Delta(USER)_{-4}$, $LVR1_{-2}$, $LVR1_{-3}$, $\Delta(USER)_{-5}$, Q1, $USER_{-4}$, $\ln(PH/PC)_{-4}$, $\ln[H/(C*OO)]_{-5}$ plus fourth order lags of the generated regressor.

Data Appendix

C - aggregate consumers' expenditure at constant prices; *UK National Accounts*.

CX - aggregate consumers' expenditure at constant prices less rents, rates and water charges; *UK National Accounts* and *Consumer Expenditure Tables*.

DUM - dummy variable equal to -1 in 1971Q2, 1 in 1972Q3 and 1988Q3 and zero elsewhere.

DUMW - dummy variable equal to (1,-1) in 1989Q4/90Q1 and (-1,1) in 1990Q3/4

D682 - dummy variable equal to 1 in 1968Q2 and zero elsewhere.

D73 - dummy variable equal to 1 in 1973Q1, -1 in 1973Q2 and zero elsewhere.

D79 - dummy variable equal to 1 in 1979Q2, -1 in 1979Q3 and zero elsewhere.

D874 - dummy variable equal to 1 in 1987Q4 and zero elsewhere.

GW - gross financial assets held by the personal sector; *Financial Statistics*.

H - stock of owner-occupied dwellings; *Housing and Construction Statistics*.

HHRATES - value of household sector payments of rates and water charges*; Consumer Expenditure Tables*, Office of National Statistics.

LVR - average loan-value ratio available to first-time buyers from building societies; *Housing and Construction Statistics.*

LVR1 = 84 - LVR

LVR2 = 78.3 + 5.7*DU - LVR, where DU is 1 from 1982Q1 onwards and zero before.

MINTQ - quarterly value of mortgage interest tax relief; *Inland Revenue Statistics*.

NL - gross liquid assets held by the personal sector less non-mortgage debt; *Financial Statistics*.

NW - net financial assets held by the personal sector; *Financial Statistics*.

OO - UK owner-occupation rate; *Housing and Construction Statistics*.

P2529 - proportion of adult population aged 25-29; *Annual Abstract of Statistics*.

PC - Consumers' expenditure deflator; *UK National Accounts*.

PCX - Implied deflator for consumers' expenditure excluding rents, rates and water charges; *UK National Accounts* and *Consumer Expenditure Tables*.

PH - DOE mix-adjusted house price index; *Economic Trends* and *Housing and Construction Statistics.*

POP - UK adult population; *Annual Abstract of Statistics*.

Q(i) - Seasonal dummy equal to unity in quarter i.

RB - clearing bank base rates; *Financial Statistics*.

REAL = RHF - $100*\Delta_4 \ln(PC)$

REP - properties taken into possession as a proportion of the number of outstanding mortgages; *Housing Finance*.

RGW = 100*GW/PC.

RHF - real post-tax cost of housing finance; *HM Treasury (1992).*

RM - average mortgage rate charged by building societies; *Financial Statistics*.

RMT = [1 - (MINTQ/(.0025*RM*SM_{-1}))]*RM; post-tax effective mortgage rate.

RNL = 100*NL/PC.

RNW = 100*NW/PC.

RS - 3 month interbank rate; *Financial Statistics*.

RUSER = [USER - 100*$\Delta_4$ln(PH)] / [1 + .01*USER - $\Delta_4$ln(PC)]

RUSERF= [USER - 100*[β*$\Delta_4$ln(PH) + (1-β)*$\Delta_4$ln(PH_{+4})]] / [1+.01*USER - $\Delta_4$ln(PC)]

RUSERX = [USER - 100*$\Delta_4$ln(PH)] / [1 + .01*USER - $\Delta_4$ln(PCX)]

RUSER1 = USER - 400*Δln(PH)

SM - stock of outstanding mortgage debt; *Financial Statistics*.

SM81 - equal to SM up to 1981Q4 and zero thereafter.

TR - basic rate of income tax; *Inland Revenue Statistics*.

USER = LVR*RMT + (1-LVR)*RB*(1-TR) + 400*HHRATES/(.01*PH*H) + 0.9

Y - real personal disposable income; *UK National Accounts*.