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Abstract

The aim of this paper is to analyse, using a vector error-correction model (VECM), the dynamic interaction between house prices and loans for house purchase in Spain. The results show that both variables are interdependent in the long run: loans for house purchase depend positively on house prices, while house prices adjust when this credit aggregate departs from the level implied by its long-run determinants. In contrast, disequilibria in house prices are corrected only through changes in this variable. As for short-run dynamics, the results show that the two variables have a positive contemporaneous impact on each other, indicating the existence of mutally reinforcing cycles in both variables.

JEL Classification: E32, G21, R21.

Key words: Mortgage Debt, Housing Prices, Error Correction.

1 Introduction

In recent years, the Spanish household balance sheet has experienced a significant change in both assets and liabilities. Indeed, household debt and, more specifically, mortgage debt, have risen rapidly, averaging 20% per annum in the period 2000-2004. In the same period, property prices have also risen significantly (16%). As a result of these changes, both household liabilities and housing wealth have greatly increased.

This considerable rise in leverage has recently attracted much attention. Indeed, although many structural factors help to explain this large increase[1], the higher level of indebtedness has raised concern, since it means a higher sensitivity of the sector to shocks to the variables that affect its debt-repayment capacity (interest rates and income, for example) and may potentially have a negative impact on spending decisions. The potential negative impact of such shocks could be especially important if they were to place simultaneously with a correction in house prices, which illustrates the relevance of the analysis of the interaction between house prices and mortgage debt and the challenges that this interaction poses for policymakers, from both a monetary policy and financial stability point of view.

From a theoretical perspective, the introduction of financial frictions as a way to explain output fluctuations has a long-standing tradition. The interaction between credit and the economy is, in fact, the essence of the financial accelerator literature[2]. Indeed, financial frictions have been introduced in theoretical models with a role of amplifying nominal or real shocks to the economy, as for example in Kiyotaki and Moore (1997), who propose a dynamic model with interaction credit limits and asset prices as a key factor behind the persistence and amplification of shocks, or Bernanke et al. (1998), where the link between the external finance premium and the net worth of potential borrowers is the key mechanism behind the propagation and amplification of shocks to the macroeconomy through credit markets. Also, many empirical papers have illustrated the significant impact that changes in net worth have on firms' and households' spending decisions (see for example Fazzari et al. (1988) for firms, and, for households, Case et al. (2001), Catte et al. (2004) or Bover (2005) for an analysis of wealth effects). In the same vein, above a certain threshold evidence of contractive effect of financial pressure on spending

[1] Namely, the liberalisation of the financial system, greater macroeconomic stability and lower financing costs derived from participation in Europe's Economic and Monetary Union, the dynamism of the labour market and the improvement of income expectations.

[2] For further details about this literature, see for example Bernanke et al. (1998)

decisions has been found in several empirical papers (see for example Hernando and Martínez-Carrascal (2003), and Martínez-Carrascal and del Rio (2004) for evidence of this contractive effect on firms' and households' spending decisions for the Spanish economy). As a result, the consideration of financial factors and asset prices in the assessment of economic prospects is a key factor for monetary policy decisions. Also, financial imbalances, asset price misalignments and the instability that may ensue when they are corrected poses important challenges to monetary policymakers, and there is in fact a growing debate about whether central banks should try to prevent the emergence of imbalances such as asset prices bubbles that can subsequently have adverse effects on economic activity (see Detken and Smets (2004) for an analysis of financial, real and monetary policy developments during and after asset price booms).

As concerns financial stability, the assessment of the soundness of the financial system obviously entails monitoring credit developments. In recent decades boom-bust cycle episodes in credit markets in some countries have resulted in financial crises and, indeed, several empirical papers have found strong evidence in favour of a positive correlation, albeit a lagged one, between credit growth and bad loans (for the Spanish case, see for example Jiménez and Saurina (2005)). In this sense, Borio and Lowe (2002) point to the credit gap as the best predictor of future problems in the financial system when compared with other indicators (equity prices and investment), and they also indicate that rapid credit growth combined with large increases in asset prices appears to increase the probability of financial instability. Moreover, house prices, not considered in this work due to a lack of adequate data, are closely linked to bank loans: they determine the value of the loans secured by property, and therefore a reduction in the value of this asset worsens financial instititutions' balance sheets and weakens their capital bases. Also, adverse housing market developments increase the proportion of non-performing loans and, as is emphasized in Hofmann (2004b), boom and bust cycles in property markets can be transmitted to credit markets and can fuel their cycles, increasing the probability of financial instability. In fact, a boom and bust in asset prices is one of the most common factors behind financial crises. Therefore, the assessment of the soundness of the financial system should take into account housing and credit market developments and, more specifically, misalignments in these variables, the interaction between both of them and their potentially self-reinforcing nature.

All these arguments illustrate the relevance of the analysis of the interaction between mortgage credit and house prices. In addition, this analysis may merit special attention for the Spanish economy for two reasons. First, the high weight of residential investment and housing wealth as a proportion of GDP and household wealth, respectively, one of the highest observed in the European Union. Second, the growth pattern of the Spanish economy in recent years, based on consumption, which has grown at a very high rate supported by the wealth effects associated with the sharp rises observed in house prices, and the dynamism of the construction sector. This growth pattern has raised concern, since it is based on housing market developments and a significant resort to borrowed funds by households and the construction and property development sectors. That poses the question of whether housing prices and indebtedness stand now at levels above those implied by their determinants. In this sense, Martínez-Pagés and Maza (2003), who model house prices as a function of gross disposable income and interest rates in the long run using an error correction model, find an overvaluation of house prices in recent years. The results of the analysis by Ayuso and Restoy (2006), who use an asset pricing model to analyze the relationship between house prices and rents, also point to the same conclusion. Martínez-Carrascal and del Rio (2004), however, do not find evidence of substantial overindebtedness in Spain in this period. However, this last result should be considered carefully because housing prices, if overvalued, may lead to a false sense of no overindebtedness. A joint model that considers endogenously both housing prices and credit may therefore be a more appropriate approach to this analysis.

This paper analyses the linkages between mortgage credit and house prices in Spain, using aggregate data. More specifically, we estimate a vector error-correction model (VECM) to test to what extent levels of house purchase debt over those implied by their long-run determinants imply adjustments in house prices, and the potential adjustment of this type of debt when house prices depart from the level implied by their determinants. Likewise, short-run dynamics are analyzed in order to determine whether there is a significant impact of contemporaneous growth of one variable on the other that results in mutually reinforcing cycles in these variables. Although a number of studies have been previously conducted to analyze the coincidence of cycles in bank credit and property prices, most of them rely on a single equation set-up (see, for example, de Haas and de Greef (2000)) and simultaneity problems imply an impossibility to draw conclusions about the direction of causality between both variables.

Our approach is in line with Hofmann (2004a) and Gerlach and Peng (2005), who analyze the interaction between bank lending and property prices for a sample of 20 countries (including Spain) and for Hong Kong, respectively. Using a VECM, they estimate a long-run cointegrating relationship for bank lending (in real terms), as a function of real GDP and real property prices, and analyze the short-run causality between both variables. Although we use the same econometric methodology, there are some relevant differences with respect to these two previous papers. First, our analysis focuses on a different credit aggregate, namely loans for house purchase (instead of bank credit), which is expected to be more linked to house prices than total bank lending (especially in those countries with a lower weight on total lending of loans backed by real estate collateral). Second, we include interest rates as a relevant determinant of this credit aggregate in the long-run. Third, we estimate a cointegrating vector for house prices, which will allow to test if loans for house purchase adjust when disequilibria in house prices are recorded. Finally, we base both our long-run and short-run analysis on the same empirical model.

Along the same lines as this paper, Fitzpatrick and McQuinn (2004) analyze the interaction of house prices and mortgage lending for the Irish economy. However, they use a single equation approach, estimating two uniequational error correction models, one for house prices and another for mortgage credit, instead of a VECM with two cointegrating relationships. Therefore, they eliminate the possibility of adjustments in one of these variables when the other registers misalignments with respect to its fundamentals.

Our results are also somewhat different from the ones found in the aforementioned papers. Both Hofmann (2004a) and Gerlach and Peng (2005) find that bank lending absorbs all the disequilibria when it departs from the level implied by its long-run determinants and, therefore, long-run causality from property prices to bank lending is obtained. Our results for Spain indicate that house prices determine the long-run level of loans for house purchase and, therefore, there is causality from the first variable to the latter one. However, we also find that house prices adjust downwards (upwards) when loans for home purchase are above (below) their long-run level, and, in this sense, our results indicate causality from the credit aggregate to house prices in the long-run disequilibria corrections. Additionally, in line with the results in Hofmann (2004a), we obtain short-run causality in both directions, while Gerlach and Peng (2005) (Fitzpatrick and McQuinn (2004)) obtain evidence of short-run causality only from house prices (mortgage credit) to bank lending (house prices) in the case of Hong Kong (Ireland).

The remainder of the paper is organized as follows. Section 2 describes the evolution of both loans for house purchase and house prices in Spain in the last two decades. Section 3 discusses the variables that can affect both variables. In Section 4, a VECM model with two cointegrating relationships – one for house prices and another for loans for house purchase – is estimated and both long-run and short-run dynamics are analyzed. Section 5 tests the robustness of the results and, finally, Section 6 summarizes the main results and concludes.

2 Loans for house purchase and house prices in Spain: some stylized facts

This section aims at describing the trend of house prices and loans for house purchase for the period covered in this analysis. Figure 1 plots the levels of these variables, together with some of their determinants. In the case of house prices, the series correspond to the price per square metre, while for loans for house purchase the logarithm of this credit aggregate is depicted.

As can be seen, house prices showed strong growth during the second half of the eighties, both in nominal and real terms. After this expansionary period, the growth rate of this variable declined substantially, and even recorded negative values. Although the decline in nominal terms was restricted to the years 1992 and 1993, the reduction of house prices in real terms lasted for longer (1992-1997) and meant a reduction of around 20%. An expansionary phase then started afresh in 1997, with especially high growth rates observed between 2000 and 2004. During these five years, house prices have more than doubled in nominal terms and the revaluation in real terms has been close to 85%.

The phases of highest growth in this variable have tended to coincide with periods of strong growth in the economy and, therefore, in household income (1986-1990 and 1997 onwards), and also with periods of high construction volumes. Indeed, both the number of dwellings per household and that of housing starts relative to the housing stock have increased during both housing booms, in an especially marked fashion in the last one (see Figure 1).

Loans for house acquisition have held on an upward trend almost throughout the period. Several factors help to explain this. On the supply side, the deregulation and liberalization of the banking system during the eighties, which resulted in better financing conditions for households; the change in the business strategy of commercial banks, which from the beginning of the eighties focused to a greater extend on the household sector, a business area that had been almost entirely dominated by savings banks until that moment; and also other factors such as the lengthening of the repayment period, which increases indebtedness capacity. Moreover, this credit aggregate was not affected by the credit restrictions in Spain between end-1989 and end-1990, which did not apply to mortgage credits. On the demand side, there was EMU entry and labour market reforms, which resulted in a reduction of financing costs, an improvement in income expectations and lower uncertainty, and involved an increase in the desired spending levels by households, and, more specifically, in housing investment. As can be seen, the growth rate of this variable is somewhat more volatile than that of house prices (although in this case no negative values have been recorded during the sample period) and has also tended to coincide with periods of robust growth in household income. The sharpest decline was observed in 1991, coinciding with the beginning of the recession of the early nineties, while the highest growth rates in real terms have been in recent years (from 1998) and also in the second half of the nineties. For the most recent period, the high growth rates of house prices have significantly increased households' indebtedness capacity, via collateral effects, and the significant rise in loans for house purchase has resulted in a substantial increase in the weight of this component relative to the liabilities of the household sector. Also, the proportion of variable-rate loans has increased over the period, accounting at present for virtually 100% of new operations, thereby increasing the sector's exposure to changes in financing costs.

Demographic factors (namely, a large number of young people becoming potential house-seekers and a sizable inflow of immigrants) have contributed to explaining the course of both variables in the latest boom cycle.

3 The determinants of loans for house purchase and house prices

There is a large body of theoretical and empirical literature on the determinants of borrowing and house prices. Theoretically, borrowing is the mechanism which, in a complete financial market environment, allows individuals to separate their spending and income flows and make their spending decisions according to their permanent income – rather than current income – and borrowing costs. Credit, therefore, adjusts passively to spending decisions. However, when there are financial frictions, as is the case in the real world, decisions cannot only be taken on the basis of permanent income and real interest rate considerations. In this case, nominal rather than real interest rates and current labour income rather than permanent income may restrict the quantity of external funds that households can obtain. Indeed, one of the key criteria banks apply in granting loans concerns the initial debt burden (namely, an initial debt burden – interest payment plus repayment of principal – lower than a given percentage of current labour income) and, consequently, a reduction in nominal interest rates, leaving real rates unchanged, can increase the quantity of debt that households can obtain (see Annex in Martínez-Carrascal and del Rio (2004) for a discussion of this point). Thus, as pointed out in Ellis (2005), this limit on the initial debt burden implies that the ratio of aggregate household debt to aggregate income converges on a long-run equilibrium level that depends, among other things, on the level of nominal interest rates.

Therefore, with credit market imperfections, current labour income, rather than permanent income, and nominal rather than real interest rates may be a better way to model borrowing to the household sector. Likewise, asymmetric information problems imply that borrowing capacity is affected by changes in house prices, which determine the collateral available for bank lending. For these reasons, in Martínez-Carrascal and del Rio (2004), labour income and nominal, instead of real, interest rates are used to jointly model borrowing and consumption. Housing wealth, found to be a key determinant of household borrowing, is also included in the specification. Also, some other recent papers, such as those of Ellis (2005) and Iacoviello (2005), have identified both nominal interest rates and collateral constraints to be relevant for mortgage debt dynamics and for the economy as a whole. Ellis (2005) analyzes the effects of income-constraint and down-payment constraints on indebtedness, while Iacoviello (2005) introduces nominal interest rates in addition to collateral constraints in a business cycle model, based on the widespread observation that in low-inflation countries most debt contracts are set in nominal terms, therefore allowing price changes to affect the realized real interest rate. This nominal characteristic of debt contracts could be especially important for the Spanish economy, where a substantial decline in the inflation rate has been observed during the sample period analyzed here. Indeed, consumers' inflation expectations in the initial stage of EMU entry were systematically above actual inflation, which implies that ex-post interest rates, based on the observed inflation rate, may not be a good proxy for ex-ante real interest rates, which are those on which investment decisions are based.

Also, the introduction of real interest rates in lending equations has empirically posed problems in previous papers that have analyzed the interaction between credit and housing prices. Gerlach and Peng (2005), who analyze the interaction between bank lending and property prices in Hong Kong, find real interest rates to be non-significant in a long-run relationship estimated for bank lending, in which they include, apart from bank lending, real GDP and real property prices. The same variables are considered by Hofmann (2004a), who omits the inclusion of real interest rates in the estimation of a cointegrating vector for this credit aggregate on the basis of to the cointegration test results for this variable. Likewise, the estimated coefficient for real interest rates in Fitzpatrick and McQuinn (2004) has a wrong (positive) sign, although quantitatively it is very small.

As for house prices, equilibrium requires the expected return, net of depreciation costs and expected capital gains or losses, to equal the return of an alternative investment with the same level of risk. This condition, together with the equilibrium condition in the market for the consumption of housing services, implies that real house prices depend on income, the housing stock and the user cost, i.e. the alternative return on investments with the same level of risk minus the expected increase in house prices net of depreciation (see Poterba (1984) for greater details of the formula derivation). More specifically, house prices are expected to depend positively on income and negatively on the user cost and the housing stock.

Empirically, however, the housing stock is often found to have a positive impact on prices, reflecting the tendency for higher increases in the number of dwellings the higher the demand for houses (and, indeed, the stock of housing in the Spanish economy has increased importantly during boom periods, as has been shown in Section 2). Likewise, an estimation of the user cost is empirically difficult because both the approximation of the return of alternative investments with the same level of risk and the estimation of the expected increase in house prices entail difficulties. This is why, very often, the user cost is proxied, albeit very roughly, by the risk-free interest rate. The high cost of housing acquisition in relation to household income also implies a close link between the housing market and the mortgage financing market. Therefore, changes in the restrictions on the supply of house financing can help to explain house price developments. The significant decline in nominal interest rates observed in Spain in recent years means that this effect is of special relevance to the Spanish case. Accordingly, in Martínez-Pagés and Maza (2003), nominal interest rates (instead of real interest rates) are used in order to analyze the evolution of house prices in Spain for the period 1977-2002.

4 Model specification and estimation results

According to the discussion in Section 3, we expect loans for house purchase and house prices to be related, in the long run, to labour income and nominal interest rates[3]. We express both credit and income in per-household terms, given that each household needs at least one house, owned or rented, for its own accommodation, and, hence, other things being equal, the higher the number of households the higher the demand for houses. All of these variables, except interest rates, are included in real terms using the private consumption deflator and expressed in log form (see the Data Appendix for a detailed definition of the variables). In addition, the difference between the return on risky assets and housing investment can have a significant impact on the behavior of house prices, given that investment can also be one of the reasons behind property acquisition. Accordingly, we add an additional short-run variable in the specification[4]: the mean difference, in the last year, between the return on mutual funds and house price increases (since in the short run, this variable may be a good proxy for housing investment return in the presence of adaptive expectations). We use mutual fund returns in order to build in a higher return than that of risk-free investments. This series that is available from 1992 to 2004. For previous years, interest rates have been used in this measure of the user cost, which seems reasonable given that household investment in shares has been very limited over this period.

Data are quarterly and cover the period 1984 Q1 to 2004 Q4. Figures 2 and 3 show the level growth and first differences of these variables. Table 1 shows the unit root tests, using a fourth lag length. According to these tests, the null hypothesis of a unit root in the levels of the series cannot generally be rejected. As for the first differences, the null hypothesis of a unit root test is always rejected when the Phillips-Perron test is used[5].

Given this assumption, equation 1 is estimated in order to determine the number of cointegrating relationships. In particular, we estimate a four-order uVAR with the constant

[3]We do not include housing stock in the specification because, in line with the results in Martínez-Pagés and Maza (2003), its inclusion posed problems with regard the signs and significance of the rest of the variables and, as in many of the previous empirical papers that model housing prices and run counter to the theory predictions, it was found to have a positive and significant impact on prices. The sensitivity of the results to changes in the variables considered will be analysed in the next section.

[4]In Section 5 the return on risky assets is also considered to be a relevant variable in the long run.

[5]When using the Augmented Dickey-Fuller test, a higher order of integration cannot be rejected for house purchase borrowing and house prices. However, since it is conceptually difficult to interpret non-stationarity of quarterly growth rates, these variables are also treated as integrated of order one.

unrestricted[6],

$$\Delta X_t = \Pi X_{t-1} + \sum_{j=1}^{q-1} \Gamma_j \Delta X_{t-j} + \delta D_t + \varepsilon_t \qquad \varepsilon_t \sim N_p(0, \Lambda) \tag{1}$$

The results[7] are shown in Table 2. As can be seen, both the trace statistic and the maximum eigenvalue test indicate, when small sample correction is used, that two cointegration relationships exist. A conditional model in which labour income is considered to be weakly exogenous, an assumption that will be tested later, also points to the existence of two cointegrating vectors (see the lower panel of the table). We then estimate a conditional model with two cointegration relationships. Also when considering the critical values proposed by Pesaran et al. (2000), that take into account the presence of I(1) exogenous variables -the user cost, in our case- the same results are obtained.

4.1 Long-run relationships

To identify the different cointegration relationships, it is necessary to impose restrictions on the β matrix. More specifically, two restrictions are needed in each equation. We impose normalization restrictions for house purchase borrowing and house prices and a unitary elasticity of borrowing to labour income[8]. Finally, in the house price equation, a zero coefficient is imposed on interest rates, the credit aggregate being the variable that captures the impact of financing costs on house prices. Hence we allow mortgage market developments not reflected in interest rate changes to have an impact on long-run house price levels. Indeed, the important changes observed in

[6] See anex B for a description of this econometric methodology.

[7] The income variable series has been adjusted for two additive outliers (Q3 1992 and Q1 1994) using TSW (TRAMO-SEATS). Three dummies have also been added in the specification. One of them is included to capture the change in the user cost computation and is given the value of 1 from year 1992 onwards. The remaining dummies capture transitory changes in loans for house purchase (1991 Q2) and return on mutual funds (2000 Q4). By their inclusion, non-normality and autocorrelation in the residuals was solved, ensuring Gaussian properties in the residuals that are necessary to obtain valid cointegrating rank tests (see Johansen and Juselius (1994) and Juselius (1994) for a discussion about this point). One additional dummy that was introduced to avoid non-normality in the labour income equation was removed from the specification when estimating the conditional model in which this variable is considered to be weakly exogenous.

[8] The validity of this identification restriction can be checked in an alternative model (Model 3) that is presented in Section 5.

the Spanish financial system as a result of its progressive liberalization and the increase in competition may have had a significant impact on house prices developments but may not be completely captured by interest rates. Moreover, the lengthening of the repayment period has increased the borrowing capacity of individuals, another factor which is not either captured by the inclusion of financing costs. In any case, the results presented here remain valid when the borrowing coefficient, rather than that associated to interest rates, is restricted to zero.

Table 3 shows the results of this exactly identified model. As can be seen, all the variables show the signs expected for both house purchase borrowing and house price equations: both variables are positively related to income, while indebtedness depends negatively on interest rates and house prices are positively related to the credit aggregate (and therefore implicitly depend negatively on financing costs). Likewise, borrowing for house acquisition depends positively on house prices, as expected. First, house prices determine the collateral available – therefore the higher the house prices, the larger the quantity of credit that households can obtain –. Second, they determine housing wealth for home owners and can therefore influence their spending and borrowing plans through positive wealth effects (the wealth effect would be negative for renters, since higher increases in property prices tend to increase housing rents, but in countries with a high degree of ownership, such as Spain, this effect is expected to be lower). All coefficients are significant at conventional significance levels, with only the p-value associated with the income coefficient in the house price equation somewhat lower than 10% (12.8%), due to the rather high standard deviation associated with the estimation of this coefficient.

Table 4 shows the results after restricting to zero those α that are not significant[9]. As can be seen, borrowing elasticity to interest rates is 10, while the elasticity to house prices is 0.78, larger than that found for total household borrowing to household wealth in Martínez-Carrascal and del Rio (2004), which is reasonable given that house purchase credit is expected to be more closely linked to house prices than total credit[10]. It is also higher than the elasticity found in Fitzpatrick and McQuinn (2004) for mortgage credit (0.51) and for total credit in Gerlach and Peng (2005) and Hofmann (2004a) for Hong Kong (0.53) and Spain (in fact, surprisingly,

[9]Although in the exacly identified model the α that captures the adjustment of loans for house purchase to house price disequilibria was significant for a 10% significance level (p-value = 8.2%), it became insignificant (p-value = 0.632) once the insignificant α associated with interest rates had been restricted to zero, while that of the coefficient of labour income in the house price equation increased (p-value = 7.2%).

[10]These elasticities are not, however, strictly comparable, since in Martínez-Carrascal and del Rio (2004) the credit aggregate is not expressed in per-household terms.

in Hofmann (2004a), and also in Hofmann (2004b), a non-significant impact of house prices on bank lending is found for Spain)[11]. Our results are therefore, more in line with those in Hofmann (2004b) for countries, such as Spain, with high home-ownership rates (those that have, consequently, higher positive potential wealth effects stemming from changes in property prices), and for those with a high weighting of real estate in total lending.

As for the house prices, they are found to depend positively on income and credit, as in Fitzpatrick and McQuinn (2004). As for the implied long-run semi-elasticity to nominal interest rates, it is found to be 4.4, very similar to that found in Martínez-Pagés and Maza (2003) when they impose a unitary elasticity of house prices to income (4.5)[12]. The elasticity of house prices to income is close to 1, and, as can be seen, the significance of this variable increases once those α that are insignificant have been restricted to zero, being the p-value associated with its estimation 8.2%.

As for the loading factors (α) that determine the dynamics of adjustment towards the long-run equilibrium, the results show that, when house purchase borrowing departs from the level implied by its determinants, the restoring of the equilibrium is achieved not only through reductions (increases) in this variable: also house prices adjust downwards (upwards) when credit is above (below) its long-run level, the speed of adjustment that results from both movements being a 5.6% per term. The adjustment in house prices (0.10 percentage point for every percentage point of disequilibrium in the credit aggregate, lower than that observed in the latter variable (0.13)) could be the result either of a decline in the desired residential investment level by households or of less willingness on the part of credit institutions to grant credit to finance housing acquisition when households are overindebted. This result is in line with the findings of Hofmann (2004a) for Denmark, Finland and Germany, while for the whole panel and also for Spain a non-significant loading factor is found, as also in Gerlach and Peng (2005) for Hong Kong. Finally, the loading factor associated with interest rates is significant, in line with the results found previously in other papers that have modelled loans using this methodology (see, for example, Calza et al. (2003)). For the Spanish case, endogeneity for interest for mortgage credit

[11] Again, these elasticities are not strictly comparable, since in Gerlach and Peng (2003), Hofmann (2004a) and Hofmann (2004b) the credit aggregate (property-related lending) is not expressed in per-household terms.

[12] Without imposing this restriction, Martínez-Pagés and Maza (2003) have problems with the significance of this variable. If we impose a zero coefficient on borrowing and a unitary elasticity to income in the housing prices equation – an specification fairly similar to theirs – the coefficient associated with interest rates rises to 7.7 and the p-value for the overidentifying restrictions is 0.70.

is found in Sastre and Fernández (2005) in which they jointly model real residential investment and house prices.

With regard to the loading factors associated with the house price cointegrating vector, it can be seen that all of the adjustment towards the equilibrium is achieved by means of reductions (increases) in this variable when it is above (below) its long-run level, with a speed of adjustment of 8% per quarter, larger in annual terms, than that obtained in Martínez-Pagés and Maza (2003) when they impose a unitary elasticity of house prices to income (28% vs 19%). Without this restriction, they obtain a rather high speed of adjustment of 49%.

Figure 4 shows the estimated long-run relationships and the error correction terms rescaled to average zero over the sample period. For the sake of comparison, the error correction term for house prices derived in Martínez-Pagés and Maza (2003) is also included. As can be seen, the disequilibrium over time of house prices with respect to their long-run level obtained with both models is similar and, according to both models, the large increase that house prices have undergone recently has placed these above their long-run equilibrium level, although to a much lesser extent than at the end of the 1980s and the beginning of the 1990s.

The overvaluation found in the last part of the sample period is, according to the model presented in this paper, somewhat lower than in Martínez-Pagés and Maza (2003) (see Figure 4). Our model, which jointly models credit and house prices, incorporates the direct impact of the credit developments on property markets, while in the Martínez-Pagés and Maza (2003) model, which uses a single-equation set-up, credit effects are captured through changes in nominal interest rates. Therefore, the lower overvaluation that we find may indicate that financial market developments other than interest rate changes may help to explain part of (but not completely) house prices evolution in the recent period. Since credit market can be in equilibrium at a given interest rate, even if supply does not equal demand (see Stiglitz and Weiss (1981)), the quantity of funds granted to the household sector can contain relevant information that is not captured by interest rates. Considering that credit restrictions have possibly diminished in the Spanish economy in recent years, it seems reasonable to think that the quantity of credit can contain relevant information, in addition to interest rates, that explain house prices developments in Spain, and a comparison of our results with those of Martínez-Pagés and Maza (2003) may point in this direction.

As for the error-correction term for borrowing for house purchase, a comparison like the one made for house prices with the results in Martínez-Pagés and Maza (2003) is not possible, since no analysis of this credit aggregate has previously been carried out for the Spanish economy. One of the closest possible comparisons for the error-correction term (ect) for the credit aggregate analyzed here would be that made using the equivalent term obtained in Martínez-Carrascal and del Rio (2004) for total household debt, in which loans for house acquisition have a predominant weighting. This comparison, illustrated in Figure 4, shows that both series show the same pattern, although the model presented here has a more pronounced upswing and downward trend, especially from 2001 onwards. Both models show that, in spite of the high growth rates of both credit aggregates during the period 2001-2003, lending error-correction terms present a downward trend, due to the considerable increase observed in house prices (and, consequently, in housing wealth in the Martínez-Carrascal and del Rio (2004) model) and the moderate reduction observed in interest rates. The downward trend is more pronounced in our model, since loans for house purchase have a higher (semi-)elasticity to both variables. The large increase in house prices observed in this part of the sample to levels over their fundamentals, together with the reduction to interest rates to historically low levels, explains why according to our model house purchase borrowing seems to be lower than the level determined by its determinants in the recent years (see Figure 6 for an illustration of the disequilibria correction. As can be seen, loans for house purchase are not under their long-run level equilibrium anymore, even under the assumption of maintenance of interest rates in the present low levels). From 2003 Q4, the ect obtained with both models shows an upward tendency, although, in this case, less pronounced in the model presented here.

As can be seen, the restrictions imposed are accepted at conventional levels (p-value = 0.82). Additionally, the weak exogeneity test assumption made for labour income can be easily accepted at conventional significance levels, as shown in Table 5.

4.2 Short-run dynamics

In this section, we turn to short-run dynamic relationships between loans for house purchase and house prices, in order to analyze the potential two-way causality between credit for property acquisition and house prices that may result in mutually reinforcing cycles in credit and property markets. More specifically, we analyze the contemporaneous interaction between both variables by estimating the corresponding equations in the system, i.e.:

$$\Delta lh_t = \alpha_{ll} CI_{lt-1} + \alpha_{pl} CI_{pt-1} + \Gamma_{pl} \Delta p_t + \Gamma_{yl} \Delta y_t + \Gamma_{il} \Delta i_t + \sum_{j=1}^{k-1} \Gamma_{jl}^{*} \Delta X_{t-j} + \Phi_l D_t + \varepsilon_{lt}$$

$$\Delta p_t = \alpha_{lp} CI_{lt-1} + \alpha_{pp} CI_{pt-1} + \Gamma_{lp} \Delta lh_t + \Gamma_{yp} \Delta y_t + \Gamma_{ip} \Delta i_t + \sum_{j=1}^{k-1} \Gamma_{jp}^{*} \Delta X_{t-j} + \Phi_p D_t + \varepsilon_{pt}$$

Following the general-to-specific approach, step by step we first remove the most insignificant variables. Additionally, in order to deal with potential simultaneity bias in the estimation of the contemporaneous coefficients estimated for Δp in the loans for house purchase equation and Δlh in the house price equation, a Hausman test is performed[13]. The test indicates that the null hypothesis of consistent ordinary least square (OLS) estimates is accepted at conventional significance levels. More specifically, the p-value associated with the residuals from the house price (loans for house purchase) equation in the loans for house purchase (house prices) equation is 0.69 (0.97).

The coefficients obtained through this estimation for contemporaneous variables and error correction terms are presented in Table 6. As can be seen, the results are consistent with those obtained in Section 4.1, since we find that lending disequilibria have a negative impact on both loans for house purchase and house prices, while lagged house price disequilibria are not significant in the loans for house acquisition equation (the magnitudes of the adjustments are, as can be seen, similar to those found in Section 4.1). In addition, the estimates for short-run parameters indicate the existence of contemporaneous short-run causality from loans for house purchase to house prices, in line with the results found in Fitzpatrick and McQuinn (2004) and in Hofmann (2004a) for his pooled estimates. The coefficient on changes in loans for house purchase in the house prices equation is positive and significant for a 10% significance level (p-value = 0.076): in the short-run housing supply is relatively rigid, given its long production process, and an increase in the external funds obtained for financing housing acquisition has a positive impact on price increases. Although positive, the coefficient estimated for contemporaneous house price growth in the loans-for-house-purchase equation is not as clearly significant. The

[13] The procedure for the Hausman test is as follows. First, an auxiliary regression for house prices (loans for house purchase) is estimated by the instrumental variable method, using the predetermined variables included in the system. Second, the equation for loans for house purchase (house prices) is re-estimated, adding the residuals from the auxiliary regression. If these residuals are not significant, the null hypothesis of consistent ordinary least square estimates cannot be rejected. Hausman tests are performed only for these two variables and not for labour income or interest rates, since we are not especially interested in the contemporaneous coefficients for these variables.

p-value associated with the estimation of this coefficient is 0.36, and therefore, according to this model, and unlike the results in Gerlach and Peng (2005) and Hofmann (2004a) for his pooled estimate and for Spain, it is not easy to reject that changes in house prices do not imply changes in the level of loans for house purchase immediately, but only with a certain delay[14]. Also, Fitzpatrick and McQuinn (2004) find that this coefficient is insignificant, although the level of significance for which this holds is not shown in their paper.

From Table 7, it can be seen that all the diagnostic tests for the residuals are passed at conventional significance levels. Likewise, one-step-ahead residuals and Chow tests, presented in Figure 7, suggest that the model is recursively stable.

5 Robustness of the results

In this section, we present the results obtained with alternative empirical specifications to check the robustness of the results presented in the previous section.

Based on the existence of credit market imperfections, the model presented in Section 4 uses nominal interest rates as the financing cost variable. As earlier pointed out, other papers, such as that of Ellis (2005), have previously illustrated the role that nominal interest rates can have in determining the level of indebtedness, while others, such us that of Iacoviello (2005), have introduced them in a business cycle model, allowing price changes to affect the realized real interest rate.

The first model in Table 7 (Model 1) shows the results obtained when real instead of nominal interest rates are used[15], once non-significant coefficients have been restricted to zero[16].

[14] Although Hofmann (2004a) finds a significant coefficient for his pooled estimate, for 15 of the 20 countries the coefficient is non-significant.

[15] The use of both real interest rates and gross disposable income, instead of labour income, resulted in mis-specifiaction problems in the borrowing equation, where real interest rates coefficient was estimated rather imprecisely and showed a positive sign. The inclusion of the inflation rate in the system as a short-run variable solved this problem, but then the house price equation showed mis-specifiaction problems.

[16] In this model, a zero coefficient is assigned to the credit aggregate in the house price equation. It may be thought that this identifying restriction in Section 4.1 above in the cointegrating vector for house prices determines the long-run direction of causality we obtain. Since loans for house purchase do not appear in the long-run relationship for house prices, it may be thought that the credit aggregate cannot adjust to its disequilibria. However, this is not the case. First, a variable that does not appear in a given estimated cointegrating vector can adjust to disequilibria (see, for example, Johansen and Juselius (1994)). In addition, the same results are

Additionally, the inflation rate has been added as a short-run variable[17], as in Nieto (2003). As can be seen, the results are substantially in line with those obtained in the model presented in Section 4. Indeed, neither the elasticities of both variables to their determinants in the long run nor the loading factors differ much from those found previously. The comparison of the error correction terms obtained with both models indicates that both series show the same course, especially for that period in which ex-post real interest rates may be a better proxy for ex-ante interest rates, that is, in the last part of the sample (see Figure 8, which shows the error correction terms obtained with both models before normalizing to zero). As earlier mentioned, during the EMU entry period consumers' inflation expectations in Spain were persistently above actual inflation, and therefore, ex-post interest rates are not a good proxy for ex-ante financing costs during this period in which the inflation rate shows a higher volatility. It is in this part of the sample where the difference between the error correction terms obtained from both models are more marked (especially in the case of the error correction term for the debt aggregate) whereas their path is quite close for the rest of the period.

Another critical point in modelling house prices is to proxy the return of alternative investments with the same risk level. As in many of the empirical papers that analyze house prices, the model presented in Section 4 uses the risk-free interest rate has been for the long-run analysis as a rough approximation, given the difficulties in obtaining a reasonable approximation for this return (in the short run, however, as previously mentioned, the difference between house price increases and risky investments has been introduced). The second and third model presented in Table 7 (Model 2 and Model 3) we check that the results presented remain valid when considering the return on risky assets in order to model house prices in the long run.

obtained when loans for house purchase explicitly enter this cointegrating vector and the interest rate coefficient is restricted to zero (in this case, a unitary elasticity to income in the house price equation is imposed, as done in Martínez-Pagés and Maza (2003). Without this restriction, the large standard deviation associated with the income coefficient, possibly due to the high correlation between the logs of this variable and the credit aggregate (95.8%), posed problems in terms of the significance of this variable. Once this unitary elasticity was imposed – a restriction accepted at conventional significance levels (p-value = 0.79) – the same results were obtained, and both income and the borrowing aggregate are significant in the house prices equation).

[17]The inflation rate, often used as a proxy for the inflation rate expectations, can affect the indebtedness decision: the higher the outlook for the inflation rate, the larger the expected reduction in the real value of the outstanding debt and, for this reason, higher inflation could stimulate the demand for new loans. On the other hand, higher inflation may be considered as an indicator of higher uncertainty and, in this sense, it could have a negative impact on the demand for new loans.

We use two different measures for alternative risky investments (see the Data Appendix for definitions)[18] with higher volatility and expected return in Model 3, which uses the IBEX 35 stock index return. Additionally, gross disposable income (instead of labour income) has been used in these specifications. Again, these estimates corroborate our previous results[19] and indicate that both house purchase credit and house prices adjust downwards (upwards) when the former is above its long-run level, although the elasticity of house prices to the return on alternative investments may be excessively large in these models[20].

As for the analysis of short-run dynamics, Table 8 shows the results obtained when the analysis performed in Section 4.2 is carried out using Model 1 in Table 7, which uses real interest rates instead of nominal ones. Again, the results indicate the existence of a contemporaneous positive impact of the growth in loans for house purchase on the growth of house prices, once potential endogeneity bias has been taken into account (as in the previous section, Hausman tests indicate that the null hypothesis of consistent OLS estimates can be accepted at conventional significance levels, being the p-value associated with the residuals from the house prices (loans-for-house-purchase) equation in the loans-for-house-acquisition (property-price) equation 0.5111 (0.514)). The lagged borrowing error-correction term is also found to be significant in the house price equation, as expected. In this model, the significance of the coefficient for contemporaneous house price growth in the equation for loans for house purchase is somewhat higher than that found in Section 4.2, but it is still low (its p-value is equal to 0.27 compared with 0.36 in Section 4.2). However, the results using Model 2 and 3 point in the same direction and, in these models, the house price growth rate is significant on the lending equation with a p-value of 5.6% and 4.3%, respectively, suggesting the existence of mutually reinforcing cycles in both variables. Again, the Hausman test indicates that consistent OLS estimates cannot be rejected in either of these two models.

Overall, the results using these alternative models corroborate the results found in Section 4 for long-run dynamics, while the short-run dynamics analysis points more clearly to the

[18]The return of risky assets is considered to be relevant only from the 1990s onwards, given that risky-asset holdings in the previous years were scarce.

[19]In Model 3, the labour income coefficient has been restricted to zero in the second cointegrating vector, because it was insignificant, even when restricting to 1 the elasticity of house prices to income. Consequently, borrowing indirectly captures the impact of both variables on house prices.

[20]Loading factors for gross disposable income and the risky-asset return considered in Model 2 do not appear in the lower panel of Table 7 because they are statistically insignificant.

existence of mutually reinforcing cycles in loans for house purchase and house prices.

6 Conclusions

This paper has analyzed the interaction between loans for house purchase and house prices in the case of the Spanish economy. First, stable long-run relationships for both variables have been estimated. Second, the process of adjustment when departures from these long-run levels are recorded has been examined. Finally, an analysis of the existing short-run dynamics between both variables has been carried out.

The analysis of long-run parameters indicates the existence of inter-dependence between both variables. On the one hand, in the long run loans for house purchase depend positively on house prices. First, the borrowing capacity of households is positively related to house prices because they determine the collateral available and, second, since house prices determine household wealth, a change in property prices can have a significant impact on spending and borrowing plans. On the other hand, however, the results show that when this credit aggregate departs from the level implied by its long-run determinants, the disequilibrium implies movements not only in this variable but also in house prices and, in this sense, there is also causality from loans for house purchase to house prices. More specifically, when credit for house purchase is above its long-run level, both variables decrease. In the case of house prices, this reduction is possibly due to either a lower desired residential investment level by households or less willingness on the part of credit institutions to grant credit when households are overindebted. In addition, the slow speed of correction of the lending disequilibrium implicit in the movements of both variables (5.6% per term) implies that the contractive impact of excessive indebtedness on house prices can be lengthy. Furthermore, given that house prices determine housing wealth, the main component of total wealth, the correction in house prices resulting from excessive borrowing for house purchase can additionally have a negative impact on consumption levels through negative wealth effects.

In contrast, house price disequilibria are fully corrected by changes in this variable (reductions when housing is overvalued and increase when housing is undervalued), and they do not imply changes in loans for house purchase. Likewise, the analysis indicates, in line with the results obtained in previous studies, that the high growth rates observed for house prices in the recent years seem to have led to an overvaluation in the price of this asset.

As for short-run dynamics, the results show that causality seems to go in both directions, since a positive contemporaneous impact of either loans for house purchase on house prices or the credit aggregate on house prices has been found, indicating the existence of mutually reinforcing cycles in both variables.

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Appendix

A Description of the variables used

lh: *Loans for house purchase per household* measured in real terms[21] (source: Banco de España)

p: *Real house prices.* The logarithm of the average price per square metre of houses (new and second-hand) (source: Ministerio de Fomento). This time series is available since 1987 Q1. Data prior to 1987 have been obtained by means of estimates of housing wealth.

i: *Nominal interest rate.* The average interest rate on bank loans to households for house purchases (source: Banco de España).

yh: *Labour income per household,* defined as the logarithm of the real labour income divided by number of households (source: INE).

r: *Real interest rate*, defined as the nominal interest rate (**i**) less inflation.

gdih: *Gross domestic income per household,* defined as the logarithm of the real gross domestic income divided by number of households (source: INE).

π**:** *Inflation* as measured by the private consumption deflator (CPI) (source: INE).

inv_r: *Risky asset return* measured as the mean real return on mutual funds (source: Banco de España) in the previous four quarters. The series goes from 1992 Q1 to 2004 Q4. For previous quarters, real interest rate has been used.

uc: *User cost of housing*, defined as the difference between **inv_r** and the mean of house price growth rate in the last four quarters.

stock_r: *Stock return*, measured by IBEX-35 log increments (source: Banco de España) from 1992.

[21] All real variables are calculated by deflating the nominal series by the CPI (source: INE).

B Econometric methodology

As stated in the introduction, a VECM will be estimated in order to analyze the interaction between loans for house acquisition and house prices. This framework involves the estimation of an unconditional qth-order vector autoregression (uVAR) over the sample $t = 1, 2, ..., T$:

$$X_t = \mu + \sum_{j=1}^{q} A_j X_{t-j} + \varepsilon_t \qquad \varepsilon_t \sim N_p(0, \Lambda), \tag{2}$$

where X_t is a vector of p variables, μ is a vector of constants, and ε_t is a p-dimensional random vector of serially uncorrelated errors with variance-covariance matrix Λ. If X_t is a cointegrated process, then the Granger representation theorem (Engle and Granger (1987)) allows VAR model to be written as a VECM:,

$$\Delta X_t = \Pi X_{t-1} + \sum_{j=1}^{q-1} \Gamma_j \Delta X_{t-j} + \delta D_t + \varepsilon_t \qquad \varepsilon_t \sim N_p(0, \Lambda) \tag{3}$$

where Δ is the first-difference operator, Γ_j are matrices of short-term parameters ($\Gamma_j = -A_{-j-1}$) and Π is a matrix of long-term coefficients ($\Pi = A_1 + A_2 - I$). We have incorporated D_t as a matrix that includes variables other than X_t that can affect ΔX_t (i.e. constants, trends, dummies, seasonal variables, or any other variables that affect the short-run behavior of variable X_t). VECM representation (3) can be used to define cointegration in a VAR system (2). A system such as that defined in 2 or 3 above is cointegrated of order r if the rank of matrix Π is equal to r. If Π is full rank ($r = p$), then X_t variables are stationary and if the rank is zero ($r = 0$), then ΔX_t is stationary or all linear combinations of X_t are $I(1)$. This rank can be obtained from the number of non-zero eigenvalues in matrix Π.

Johansen (1991) defined tests on the rank of Π based on a general Likelihood Ratio (LR) test. Under the null hypothesis, the cointegration rank is at most r ($H(r)$). Johansen (1991) proposed two alternative hypotheses. The first (H_1) implies that Π is full rank ($r = p$). The LR test statistic is given by equation 4 and is called the trace statistic:

$$\tau_{trace} = -2 \ln(Q; H(r)|H_1) = -T \sum_{i=r+1}^{p} \ln(1 - \hat{\lambda}_i) \tag{4}$$

where $\hat{\lambda}_i$ are the estimated eigenvalues of matrix Π. Alternatively, Johansen (1991) proposed a LR test where the alternative is that rank is equal to $r + 1$ ($H(r + 1)$). In this case the LR test statistic adopts the form of equation 5 and is called the maximum eigenvalue statistic.

$$\tau_{\lambda_{max}} = -2\ln(Q; H(r)|H(r+1)) = -T\ln(1-\hat{\lambda}_{r+1}) \tag{5}$$

Once a suitable number of cointegration relationships (r) has been selected, it is possible to decompose matrix Π into a product of two matrices ($\Pi = \alpha\beta$), where α and β are $(p \times r)$ and $(r \times p)$ (full rank) matrices of loading factors and long-run coefficients, respectively. Even though the number of cointegrating relationships can be determined, vectors α and β cannot be estimated, since there is an infinite number of matrices that satisfies $\Pi = \alpha\beta$. Some restrictions (at least r^2), derived from economic theory, should be made in order to identify loading factors and long-run relationships.

Sometimes we are only interested in the stochastic properties of some of the variables of vector X_t. Let us define Y_t as the vector of the g variables in which we are interested, and Z_t as the rest of the variables $(p-g)$ of vector X_t. The use of a VECM allows the breakdown of system 3 and the formulation of a partial system on Y_t as a conditional model (6):

$$\Delta Y_t = \alpha^*\beta X_{t-1} + \sum_{j=1}^{q-1}\Gamma_j^*\Delta X_{t-j} + \Phi\Delta Z_t + \delta^* D_t + \varepsilon_{yt} \qquad \varepsilon_{yt} \sim N_g(0,\Sigma^*) \tag{6}$$

where $\alpha^* = \alpha_y - \Phi\alpha_z$, $\Gamma_j^* = \Gamma_{jy} - \Phi\Gamma_{jz}$, $\delta^* = \delta_y - \Phi\delta_z$, $\Phi = \Lambda_{yz}\Lambda_{zz}^{-1}$ and $\Sigma^* = \Sigma_{yy} - \Sigma_{12}\Sigma_{zz}^{-1}\Sigma_{21}$. Variables included in Z_t are explained by marginal model 7:

$$\Delta Z_t = \sum_{j=1}^{g-1}\Gamma_{j2}\Delta X_{t-j} + \alpha_2\beta X_{t-1} + \delta_z D_t + \varepsilon_{zt} \qquad \varepsilon_{zt} \sim N_{p-g}(0,\Sigma_{zz}) \tag{7}$$

All the βX_{t-1}cointegrating relations enter into both marginal and conditional models, and that new loading factors (α^*) are dependent on the loading factors of Z (α_z). In the general case, parameters of both model 6 and model 7 are interrelated, which means that full system analysis is needed to draw efficient inference about parameters (Johansen (1992b)).

The special case in which conditional model 6 contains as much information as the full system about loading factors and long-run coefficients, vector Z_t is referred as being *weakly exogenous* for α and β (Engle et al. (1983)). The necessary and sufficient condition for long-run weak exogeneity of Z_t is that the loading factors of the long-run relationship (α_z) are zero. Models 6 and 7 can therefore be reduced to the following expressions:

$$\Delta Y_t = \alpha_y \beta X_{t-1} + \sum_{j=1}^{k-1} \Gamma_j^* \Delta X_{t-j} + \Phi \Delta Z_t + \delta^* D_t + \varepsilon_{yt} \qquad \varepsilon_{yt} \sim N_g(0, \Sigma^*) \tag{8}$$

$$\Delta Z_t = \sum_{j=1}^{k-1} \Gamma_{j2} \Delta X_{t-j} + \delta_z D_t + \varepsilon_{zt} \qquad \varepsilon_{zt} \sim N_{p-g}(0, \Sigma_{zz}) \tag{9}$$

In this case, β and α_y parameters are present only in equation 8, and parameters in both models are variation-free (see Johansen (1992a) and Urbain (1992)). Vector ΔZ_t still depends on ΔY_{t-i} in equation 9[22]. Model 8 can be estimated independently of model 9 and has a small number of parameters in comparison with joint model 3. Model 8 will be the model used in this paper.

[22] If coefficients associated with ΔY_{t-i} are equal to zero, then Y_t does not Granger-cause ΔZ_t and is strongly exogenous for β.

Table 1. Unit root tests

		level				1st difference			
		ADF		PP		ADF		PP	
lh_t	None	3.72		14.07		-0.19		-4.38	***
	Intercept	2.26		4.28		-2.02		-9.10	***
	Trend	-1.91		-2.22		-3.21	*	-11.14	***
p_t	None	1.30		3.44		-1.08		-3.50	***
	Intercept	-1.35		-0.05		-1.70		-4.90	***
	Trend	-2.91		-0.88		-1.69		-4.90	***
i_t	None	-1.51		-1.79	*	-3.94	***	-3.84	***
	Intercept	-0.09		-0.29		-4.36	***	-4.04	***
	Trend	-1.81		-1.67		-4.49	***	-4.03	**
yh_t	None	1.84		1.58		-3.61	***	-11.92	***
	Intercept	-1.11		-0.55		-4.16	***	-12.45	***
	Trend	-2.03		-2.72		-4.15	***	-12.37	***
r_t	None	-0.75		-0.84		-5.51	***	-10.02	***
	Intercept	-0.64		-1.17		-5.53	***	-9.98	***
	Trend	-2.30		-2.45		-5.88	***	-10.13	***
$gdih_t$	None	2.77		2.50		-3.85	***	-11.63	***
	Intercept	-1.25		-0.72		-4.84	***	-12.78	***
	Trend	-1.94		-2.81		-4.86	***	-12.73	***
π_t	None	-2.17	**	-2.54	**	-5.22	***	-12.49	***
	Intercept	-2.58		-2.90	**	-5.38	***	-12.71	***
	Trend	-3.09		-3.57	**	-5.53	***	-12.93	***
uc_t	None	-0.64		-1.27		-2.42	**	-4.03	***
	Intercept	-0.73		-1.36		-2.42		-4.03	***
	Trend	-1.00		-1.50		-2.40		-4.01	**
inv_r_t	None	-1.41		-2.11	**	-5.85	***	-9.90	***
	Intercept	-1.70		-2.61	*	-5.81	***	-9.84	***
	Trend	-2.36		-3.14		-5.79	***	-9.79	***
$stock\_r_t$	None	-2.13	**	-3.40	***	-6.04	***	-10.37	***
	Intercept	-2.24		-3.46	**	-5.99	***	-10.28	***
	Trend	-2.22		-3.46	*	-5.94	***	-10.20	***

Note: The null hypothesis is the presence of a unit root and ***, ** and * denotes the rejection of the null hypothesis at significance levels of 1%, 5% and 10% respectively, based on critical values by MacKinnon (1994). ADF denotes the Augmented Dickey Fuller test (with lags up to and including the highest lag statistically significant at least at the 5% level). PP denotes the Phillips Perron test.
lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household; r_t real interest rate; $gdih_t$ gross domestic income per household; π_t inflation; inv_r_t risky asset return; uc_t user cost of housing; stock_r Stock return (see Data Appendix for a description of the variables).

Table 2. Johansen tests for cointegration

Unrestricted model

H_o: rank=r	λ-max test statistic	Small sample correction	95% critical values	95% Pesaran *et al* (2000) critical values	Trace test statistic	Small sample correction	95% critical values	95% Pesaran *et al* (2000) critical values
$r = 0$	64.26**	51.41**	27.1	30.7	115.3**	92.22**	47.2	58.6
$r \leq 1$	32.22**	25.78**	21	24.6	51.01**	40.81**	29.7	38.9
$r \leq 2$	**17.52***	**14.02**	**14.1**	**18.06**	**18.79***	**15.03**	**15.4**	**23.3**
$r \leq 3$	1.274	1.019	3.8	11.47	1.274	1.019	3.8	11.5

Diagnostic tests

	LM test		Normality test		Arch test		Heteroscedasticity	
Single equation tests								
lh_t	0.453	[0.809]	0.039	[0.981]	1.476	[0.226]	0.871	[0.645]
p_t	0.410	[0.840]	2.743	[0.254]	0.329	[0.857]	0.419	[0.985]
i_t	1.584	[0.183]	3.740	[0.154]	1.098	[0.370]	0.610	[0.894]
yh_t	1.174	[0.213]	5.625	[0.060]	0.994	[0.421]	0.675	[0.841]
System			10.859	[0.210]			0.515	[1.000]

Number of lags used in the analysis: 4

Variables entered unrestricted: uc_{t-1} uc_{t-2} uc_{t-3} uc_{t-4}
Constant $CSeason_1$ $CSeason_2$ CSeason

Restricted model

Variables entered restricted: labour income

H_o: rank=r	λ -max test statistic	Small sample correction	95% critical values	95% Pesaran *et al.* (2000) critical values	Trace test statistic	Small sample correction	95% critical values	95% Pesaran *et al.* (2000) critical values
$r = 0$	59.24**	50.36**	21	27.75	95.01**	80.76**	29.7	46.4
$r \leq 1$	34.26**	29.12**	14.1	21.07	35.77**	30.41**	15.4	28.4
$r \leq 2$	**1.51**	**1.29**	**3.80**	**14.35**	**1.51**	**1.29**	**3.80**	**14.35**

Diagnostic tests

	LM test		Normality test		Arch test		Heteroscedasticity	
Single equation tests								
lh_t	0.673	[0.646]	0.626	[0.731]	0.919	[0.461]	1.119	[0.391]
p_t	0.307	[0.906]	2.541	[0.281]	0.145	[0.964]	0.528	[0.944]
i_t	1.298	[0.281]	3.503	[0.174]	0.942	[0.449]	0.741	[0.774]
System	1.085	[0.360]	6.869	[0.333]			0.762	[0.946]

Number of lags used in the analysis: 4

Variables entered unrestricted: uc_{t-1} uc_{t-2} uc_{t-3} uc_{t-4} Δyh_t Δyh_{t-1} Δyh_{t-2} Δyh_{t-3}
Constant $CSeason_1$ $CSeason_2$ CSeason

Variables entered restricted: yh_t

Note: The VAR model includes four lags of endogenous variables specified in levels. The constant term is unrestricted. * (**) denotes the existence of cointegration at a significance level of 5% (1%). In the diagnostic tests, p-values are in brackets. The LM test is the Godfrey test for auto-correlation. For normality the Doornik and Hansen test is used. The Arch test is for the auto-regressive conditional heteroscedasticity test. The heterocedasticity test is a White test for individual equations while the Doornik and Hendry test is for the whole system. See Doornik and Hendry (1998) for more details. Pesaran et al (2000) critical values take into account the inclusion of I(1) exogenous variables.
lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household (see Data Appendix for a description of the variables).

Table 3. Conditional system. Exactly identified cointegrating vectors

Long-run coefficient

	lh_t	p_t	i_t	yh_t
β_1	**1**	**-0.843**	**9.28**	**-1**
std errors		0.084	0.578	0
β_2	**-0.359**	**1**	**0**	**-1.037**
std errors	0.134		0	0.682

Loading factors

	Disequilibrium in			
	Vector 1 lh		Vector 2 p	
	α	std errors	α	std errors
Δlh_t	**-0.145**	0.028	-0.027	0.016
Δp_t	**-0.104**	0.027	**-0.094**	0.015
Δi_t	**-0.029**	0.006	-0.005	0.003

Note: lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household (see Data Appendix for a description of the variables).

Table 4. Conditional system. Over-identified cointegrating vectors

Long-run coefficient

	lh_t	p_t	i_t	yh_t
β_1	1	**-0.784**	**10.05**	-1
std errors	0	0.079	0.523	0
β_2	**-0.328**	**1**	0	-1.200
std errors	0.132		0	0.691

Loading factors

	Disequilibrium in			
	Vector 1 lh		Vector 2 p	
	α	std errors	α	std errors
Δlh_t	**-0.132**	0.025	-	-
Δp_t	**-0.096**	0.025	**-0.080**	0.014
Δi_t	**-0.027**	0.005	-	-

Overidentifying restrictions test

LR-test, rank = 2: $\chi^2(2) = 0.387$ [0.824]

Note: lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household (see Data Appendix for a description of the variables).

Table 5. Exogeneity tests

Weak exogeneity for long-run parameters

Variables in marginal model: yh_t

Regressor	Distribution:			
ECT_lh_{t-1}	F(1,56)	=	0.44	[0.5076]
ECT_p_{t-1}	F(1,56)	=	1.26	[0.2672]
ECT_lh_{t-1}, ECT_p_{t-1}	F(2,56)	=	0.83	[0.4404]

Note: lh_t denotes loans for house purchase per household; p_t, real housing prices; yh_t labour income per household (see Data Appendix for a description of the variables); ECT denotes error correction term (associated with the overidentified cointegrating vectors).

Table 6. Short-run dynamics

	Δlh_t	Δp_t
Δlh_t	-	0.150 (0.083)*
Δp_t	0.091 (0.099)	-
Δi_t	-	-
Δyh_t	0.225 (0.013)*	-
$CI(lh)_{t-1}$	-0.125 (0.021)***	-0.079 (0.011)***
$CI(p)_{t-1}$	-	-0.069 (0.010)***
R^2	57.4%	74.6%

	Residual tests			
LM test	0.546	[0.741]	0.166	[0.974]
Normality test	0.656	[0.720]	2.660	[0.265]
Arch test	0.642	[0.635]	0.893	[0.474]
Heteroscedasticity	2.059	[0.056]	1.112	[0.367]

Note: Figures in parenthesis represent standard deviations. *, ** and *** denotes significance levels of 10%, 5% and 1% respectively.
lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household (see Data Appendix for a description of the variables).

Table 7. Alternative models. Long-run analysis

	Model 1				**Model 2**				**Model 3**			
	Cointegrating vector for lh		**Cointegrating vector for p**		**Cointegrating vector for lh**		**Cointegrating vector for p**		**Cointegrating vector for lh**		**Cointegrating vector for p**	
	β_1	std errors	β_2	std errors	β_1	std errors	β_2	std errors	β_1	std errors	β_2	std errors
lh_t	1		0	0	1		0	0	1		-0.483	0.146
p_t	-0.772	0.083	1	0	-0.876	0.057	1	0	-0.907	0.098	1	0
i_t					10.756	0.424	0	0	10.935	0.545	0	0
yh_t	-1	0	-1.770	0.929						0	0	0
r_t	10.104	0.547	3.919	1.570								
$gdih_t$					-1	0	-2.518	0.958	-0.7784	0.481		
inv_r_t					0	0	5.096	0.648				
$stock\_r_t$									0	0	1.122	0.154

	Loading factors				**Loading factors**				**Loading factors**			
	Vector 1 lh		**Vector 2 p**		**Vector 1 lh**		**Vector 2 p**		**Vector 1 lh**		**Vector 2 p**	
	α_1	std errors	α_2	std errors	α_1	std errors	α_2	std errors	α_1	std errors	α_2	std errors
Δlh_t	-0.123	0.026	-	-	-0.134	0.021	-	-	-0.127	0.019	-	-
Δp_t	-0.053	0.022	-0.046	0.009	-0.134	0.023	-0.047	0.007	-0.128	0.024	-0.047	0.008
Δi_t					-0.022	0.004	-0.004	0.001	-0.024	0.004	-0.004	0.002
Δr_t	-0.029	0.006	-	-								
$\Delta stock\_r_t$							-0.030	0.013	-	-	-0.242	0.078

	Overidentifying restriction test. P-value	**Overidentifying restriction test. P-value**	**Overidentifying restriction test. P-value**
	0.743	0.711	0.554

	Weak exogeneity test for labour income. P-value	**Weak exogeneity test for labour income. P-value**	**Weak exogeneity test for labour income. P-value**
Vector 1	0.47	0.78	0.72
Vector 2	0.54	0.54	0.53
Vector 1 and 2	0.60	0.74	0.72

Number of lags used in the analysis: 4

Variables entered unrestricted:

Model 1	Model 2	Model 3
$CSeason_1$ $CSeason_2$ CSeason	$CSeason_1$ $CSeason_2$ CSeason	$CSeason_1$ $CSeason_2$ CSeason
Constant uc_{t-1} … uc_{t-4} π_{t-1} … π_{t-4}	Constant	Constant

Note: lh_t denotes loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household; r_t real interest rate; $gdih_t$ gross domestic income per household; π_t inflation; inv_r_t risky asset return; uc_t user cost of housing; stock_r stock return (see Data Appendix for a description of the variables).

Table 8. Alternative models. Short-run dynamics.

	Model 1		Model 2		Model 3	
	Δlh_t	Δp_t	Δlh_t	Δp_t	Δlh_t	Δp_t
Δlh_t	-	0.153 (0.079)*	-	0.203 (0.089)**	-	0.336 (0.087)***
Δp_t	0.100 (0.09)	-	0.158 (0.088)*	-	0.170 (0.082)**	-
Δr_t	-	-	-	-	-	-
Δyh_t	-	-	0.362 (0.090)***	-	0.446 (0.098)***	-
$\Delta\pi_{t-1}$	-	-0.715 (0.127)***	-0.108 (0.013)***	-0.109 (0.021)***	-0.106 (0.011)***	-0.061 (0.009)***
$CI(lh)_{t-1}$	-0.118 (0.024)***	-0.057 (0.007)***	-	-0.042 (0.006)***	-	-0.045 (0.008)***
$CI(p)_{t-1}$	-	-0.036 (0.007)***	-	0.203 (0.089)**	-	0.336 (0.087)***
R^2 Adjusted	58.9%	82.8%	63.5%	70.8%	67.6%	70.2%
Residual tests						
LM test	0.742 [0.595]	0.377 [0.862]	0.515 [0.764]	0.820 [0.540]	0.779 [0.569]	0.380 [0.861]
Normality test	0.159 [0.924]	1.675 [0.433]	8.839 [0.012]	1.251 [0.535]	7.427 [0.024]	7.861 [0.020]
Arch test	0.483 [0.748]	0.307 [0.872]	0.280 [0.890]	1.523 [0.208]	0.283 [0.888]	0.026 [0.999]
Heteroscedasticity	1.091 [0.390]	0.491 [0.954]	0.928 [0.551]	1.122 [0.366]	1.831 [0.054]	0.327 [0.994]

Note: Figures in parenthesis represent standard deviations*, ** and *** denotes significance levels of 10%, 5% and 1% respectively.
CI(·) denotes the restricted cointegration vectors obtained in the long-run analysis of each model; lh_t loans for house purchase per household; p_t, real housing prices; i_t nominal interest rate; yh_t labour income per household; r_t real interest rate; $gdih_t$ gross domestic income per household (see Data Appendix for a description of the variables)

Figure 1. loans for house purchase, house prices and their determinants. Evolution over time

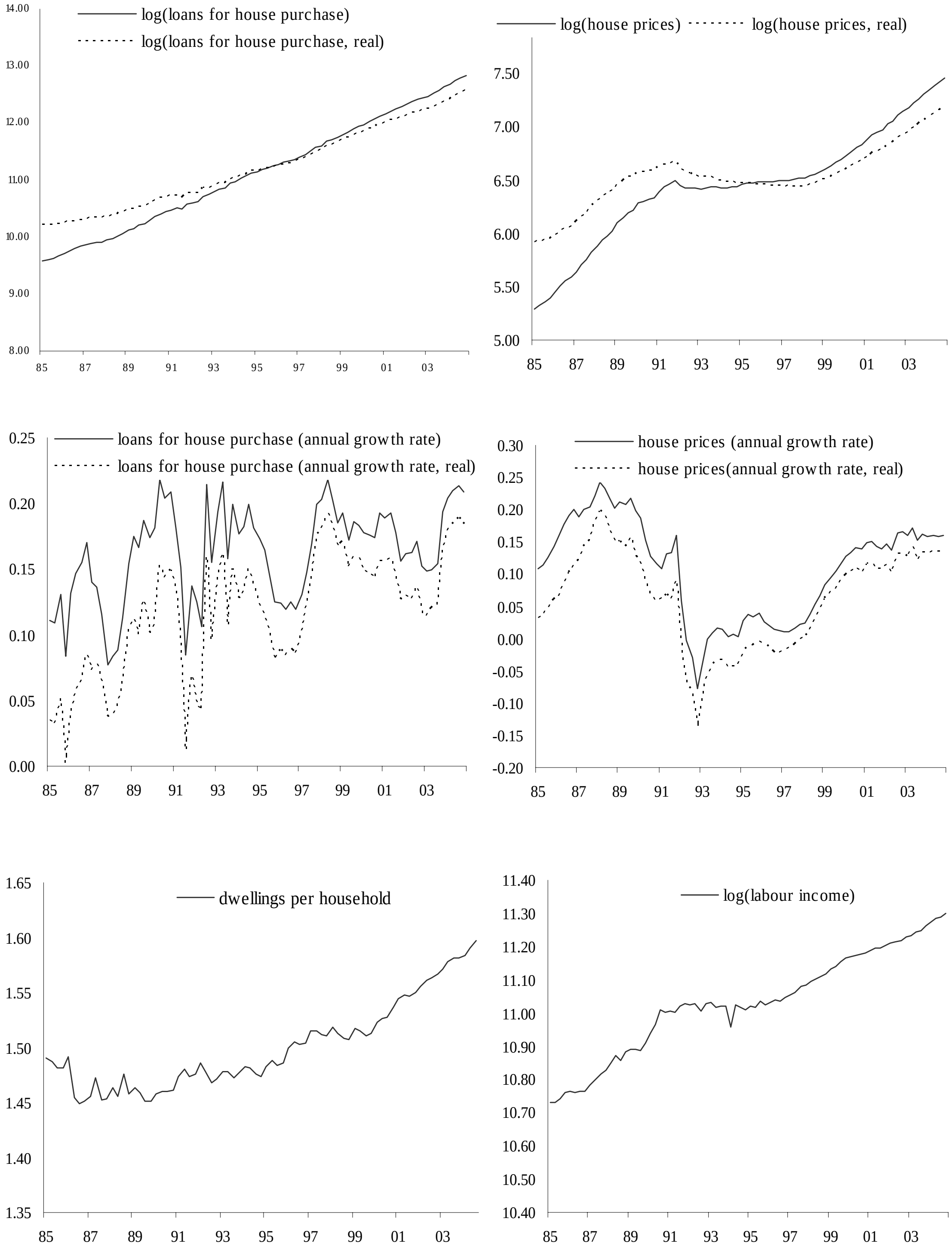


Figure 2. variables included in the models

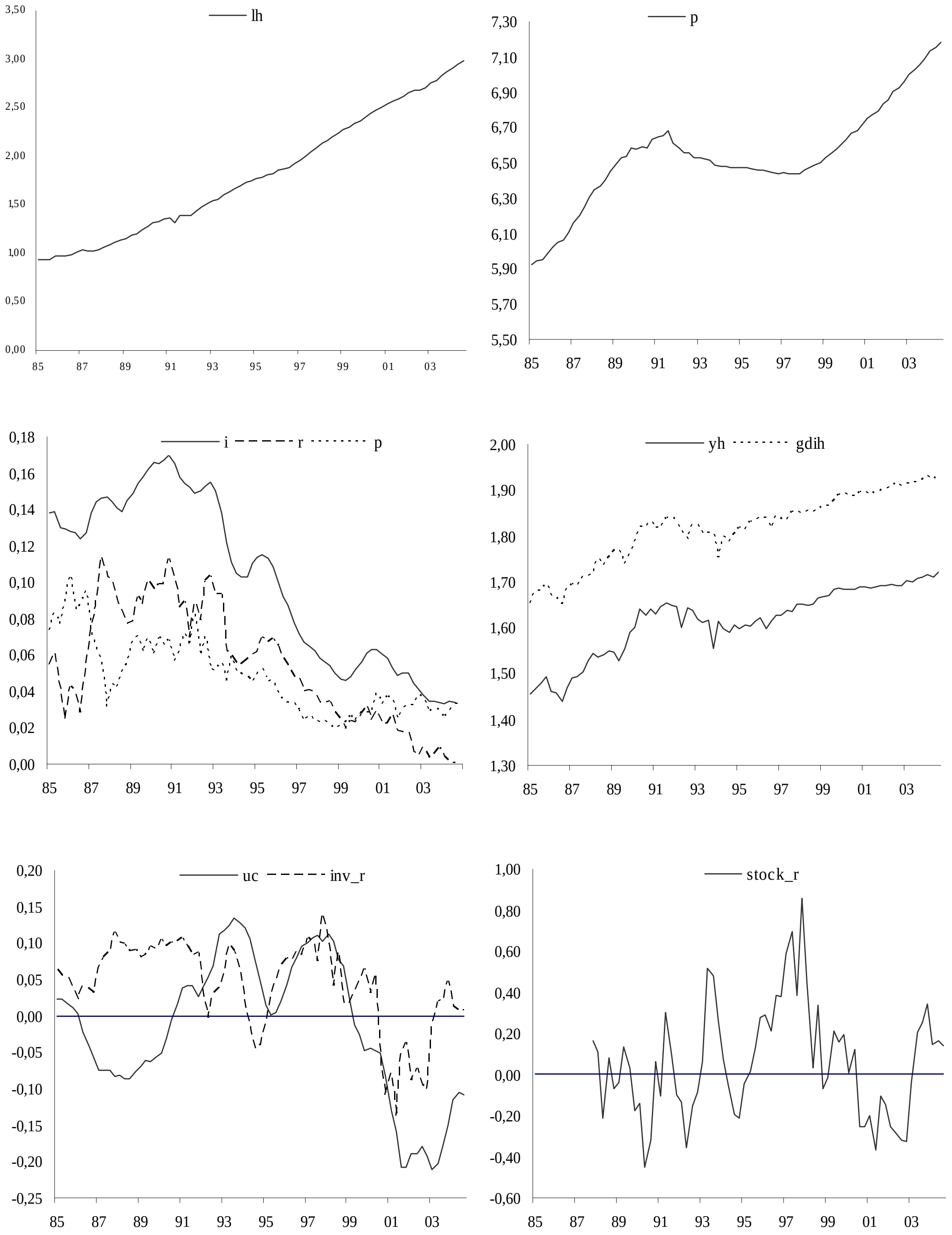


Figure 3. First differences of explanatory variables

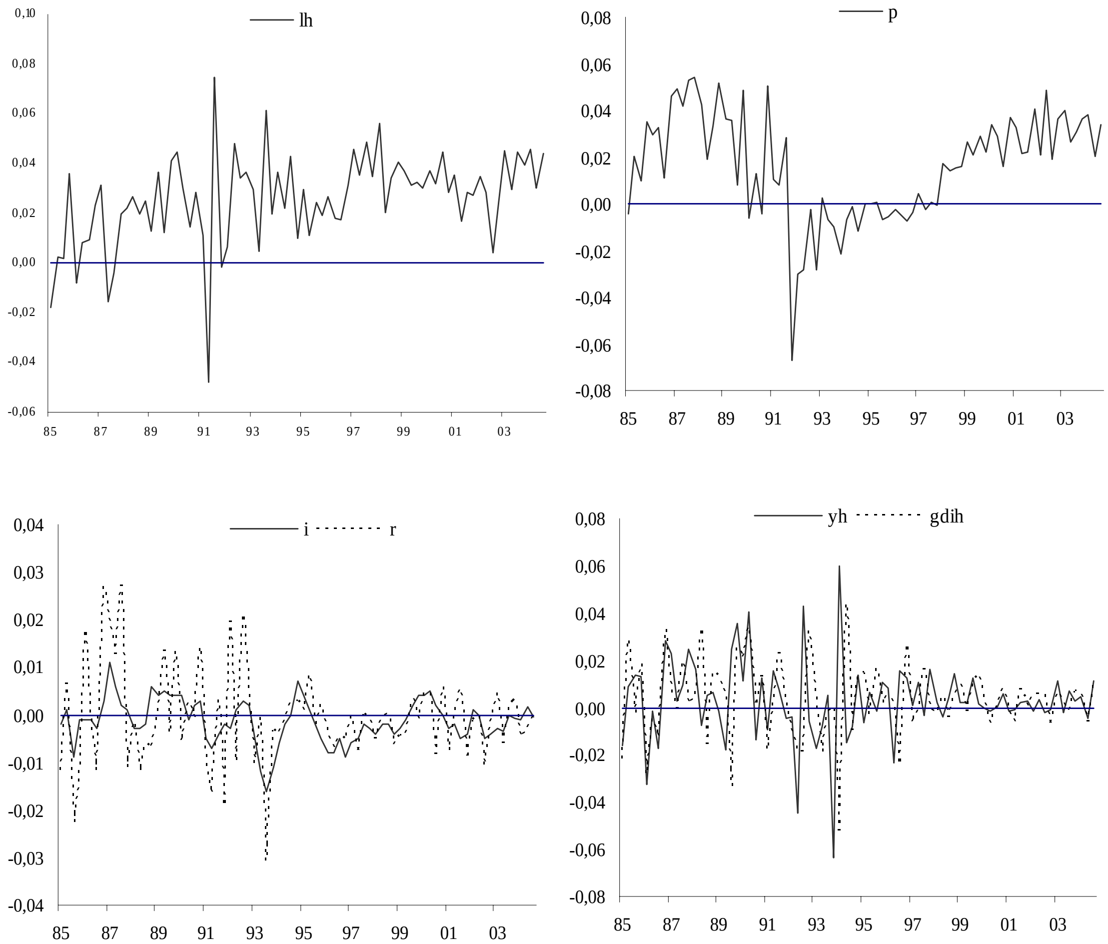
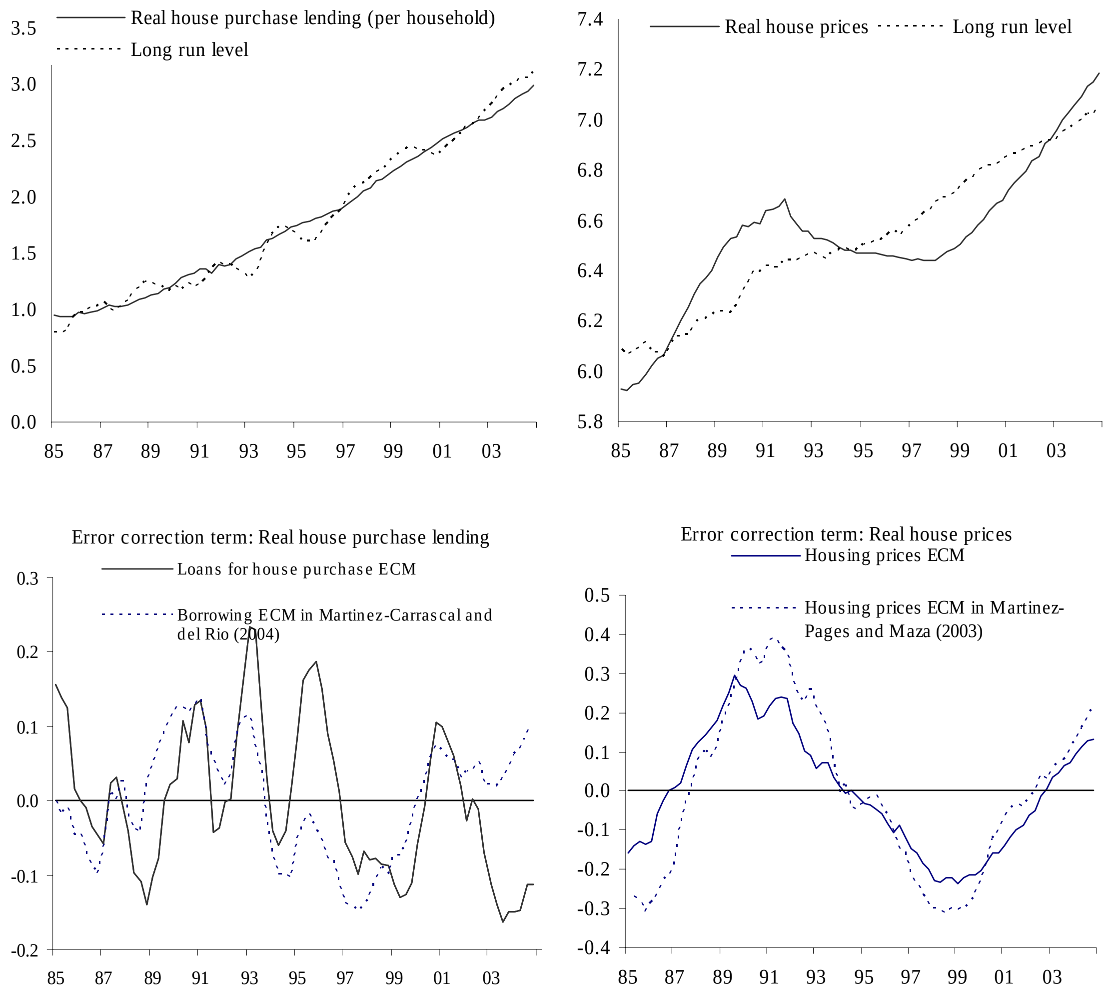


Figure 4. Estimated long-run relationships and error-correction terms



Note: Re-scaled to average zero.

Figure 5. Fitted variables and residuals

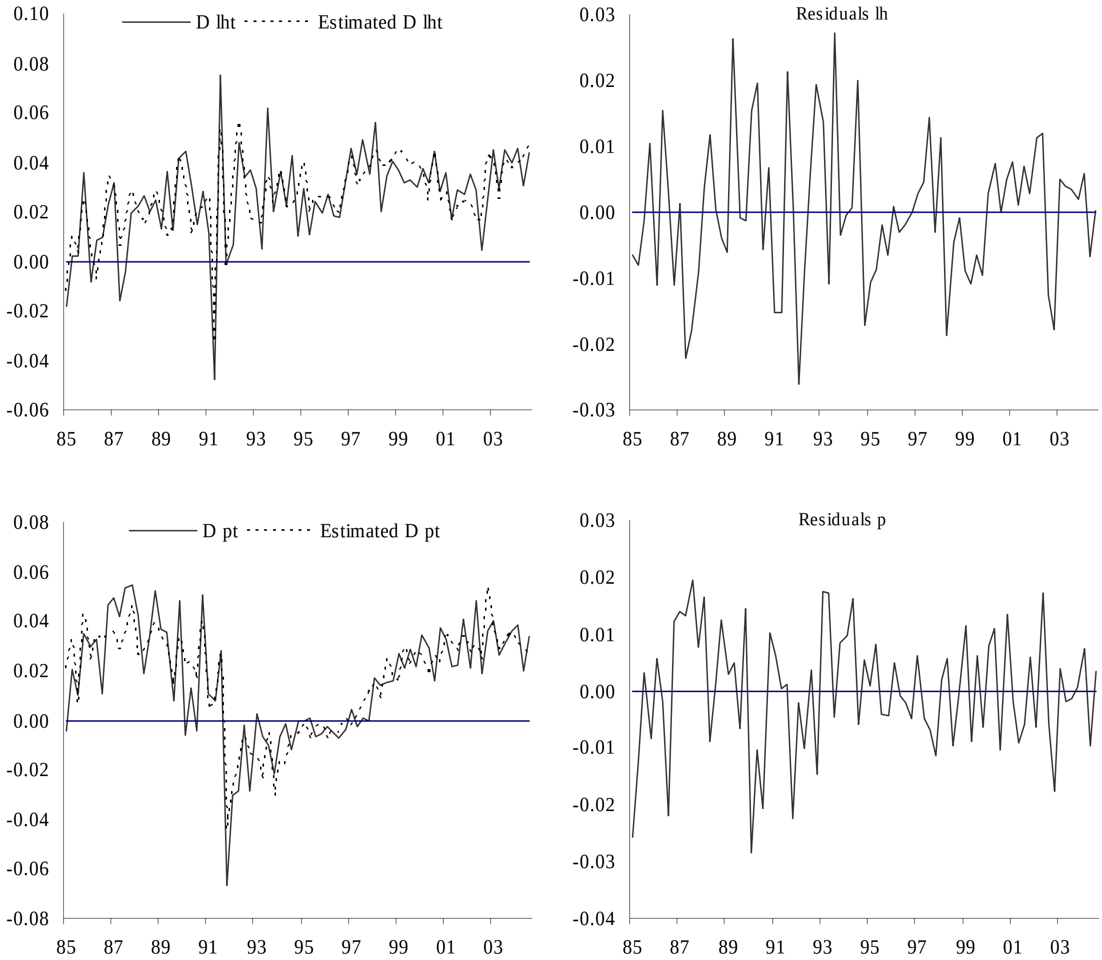


Figure 6: Loans for house purchase and housing prices. Simulated transitions to equilibrium

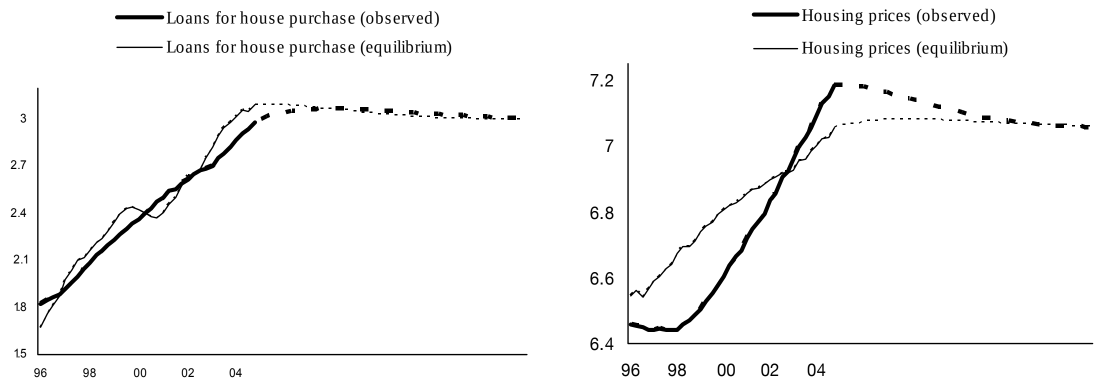
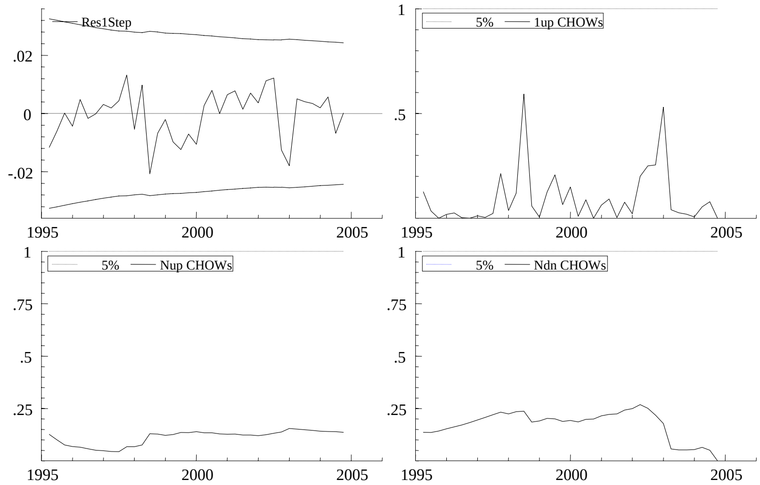


Figure 7. One-step-ahead residuals and Chow tests

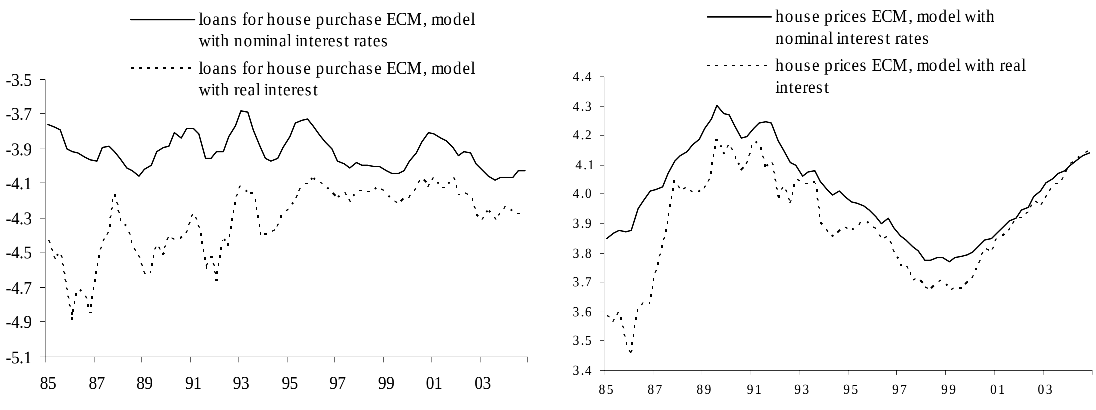
Lending equation



House price equation

Note: Panel 1 are one-step ahead residuals for each equation; Panel 2 (top right) one-step Chow tests. Panel 3 (botton left) forecast Chow test and Panel 4 break point Chow tests.

Figure 8: Nominal vs real interest rates. Error correction terms comparison



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