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Self-reinforcing effects between housing prices and credit: Evidence from Norway

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Self-reinforcing effects between housing prices and credit: Evidence from Norway*

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Abstract

The interaction between housing prices and household borrowing in Norway is estimated in a simultaneous setting in the long and the short run. The long run dependence is analyzed within a cointegrated vector autoregression in real housing prices, real disposable household income and real household debt, conditioning on the real after tax interest rate, the number of house transactions and the volume of housing capital. We identify two cointegrating equations which determine equilibrium housing prices and household debt, respectively. The long run equations are embedded in a system of two error-correction equations which is estimated simultaneously. The model yields meaningful short and long term effects when estimated on the sample 1986q2-2008q4 and impulse responses demonstrate that there are self reinforcing feedback effects between the two variables of interest.

Keywords: *Housing prices, household borrowing, financial accelerator*

JEL classification: *C32, C52, E44, G21, G28*

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*This paper builds on Anundsen (2010) and it is a slightly revised version of Anundsen and Jansen (2011).

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1 Introduction

The world wide financial crisis that originated with the US sub-prime crisis of 2007 has highlighted the importance of the interplay between financial markets and the real economy. A great number of factors contributed to the current crisis [see IMF (2009), Hubbard and Mayer (2009) and Acharia and Schnabl (2009)]. However, it seems to be widely agreed that it was primarily an unsustainable weakening of credit standards that induced the US mortgage lending and housing bubble. Other countries with more stable credit conditions were mainly affected through the international financial linkages, *e.g.* European banks incurring heavy losses on securities tightly connected to the US mortgage market in the wake of the meltdown. In those countries, as Duca et al. (2010) emphasize, any overshooting of construction and housing prices owed more to traditional housing supply and demand factors.

However, there is a two-way direction of causation since imbalances in the housing market, or the real estate market in general, oftentimes have threatened the stability of the financial sector. In the past there have been numerous episodes where falling housing prices have preceded financial crises, as Koetter and Poghosyan (2010) point out. They also argue that, due to decentralized trading with imperfect information and high transaction costs on the one hand and slow supply responses due to construction lags and limited land availability on the other, sustained deviations from long run equilibria occur more frequently in the housing market than in the financial markets.

In the real estate market and the housing market in particular the amount of credit made available by lenders depends on the net worth of the debtors. Due to imperfections and informational asymmetries in the credit markets, a prospective borrower is usually granted a loan only by putting up collateral. In the models developed by Kiyotaki and Moore (1997) and Bernanke and Gertler (1989) shocks to the real economy are amplified through the credit market by altering the value of borrowers' net worth.

This so-called *financial accelerator*[1] mechanism offers an explanation to the housing market fluctuations. First, higher housing prices increase the amount of credit needed to finance a given housing purchase. Thus, we would expect higher property valuations to put an upward pressure on the demand for credit. Second, most housing loans are secured by the property itself. An increase in housing prices raises the value of the housing capital, which feeds into a greater net worth for the household sector. By increasing the net worth and thus the value of the collateral, higher housing prices will increase their borrowing capacity. At the same time, higher property valuations make banks' assets less risky, as the increased value of the collateral pledged reduces the likelihood of defaults on existing loans, which may stimulate the banks to expand their lending.

On the other hand, most housing purchases are financed by credit, and changes in household borrowing are expected to affect housing prices through liquidity effects. The potential self-reinforcing mechanism that works across these markets makes it important to study from the perspective of financial stability, and it constitutes a main reason why central banks commonly assess financial sector vulnerability by monitoring both property prices and credit growth. The close relationship between

[1] The term was coined in Bernanke and Gertler (1995), see also Bernanke et al. (1999).

the evolution of property prices and credit aggregates has been a focal point in the policy-oriented literature, see *e.g.* Borio et al. (1994).

In this paper we pay heed to the interaction between Norwegian housing prices and household debt by analyzing the two in a simultaneous setting. First, in Section 2 we give a brief survey of recent literature. A description of the Norwegian housing and credit markets is outlined in Section 3. In Section 4 we investigate the fundamental determinants of housing prices and household debt in a multivariate cointegration framework motivated by a simple theory model. Section 5 describes the dynamic interaction between the two variables. The model yields meaningful short and long term effects when estimated on the sample 1986q2-2008q4. An interesting feature of the model is that an expectations variable has a significant short term effect in the determination of housing prices. In Section 6 we report impulse responses which demonstrate that there are self-reinforcing feedback effects between the two variables of interest. Section 7 concludes.

2 A survey of recent empirical contributions

A nascent literature on the interaction between housing prices and credit has recently emerged on the international arena. In an early study Hofmann (2003) finds that changes in property prices affect private sector borrowing in the long run when analyzing a panel comprising quarterly data for 20 countries (Norway included) covering the period 1985-2001. The opposite direction of causation is not supported. The author also reports results for the short run adjustment, where he finds that the causality goes in both directions. The long run results are further corroborated in Hofmann (2004)[2] where separate VARs in real credit to the private sector, GDP (as a broad measure of economic activity) and the short-term real interest rate as a measure of financing costs are first studied for each country. For a clear majority of the countries Johansen analysis shows no cointegration with this information set. When extending the analysis to include real property prices in the VARs, Hofmann finds strong support for one cointegrating vector for all countries, which (through the significance of the loadings) can be interpreted as determining credit in the long run for those countries where a high share of loans are secured by real estate.

These results are supported by Brissimis and Vlassopoulos (2009) in a single country study for Greece, for quarterly data specific to the housing market covering the period 1993-2005. Using multivariate cointegration techniques they find, based on a cointegrated relationship identified as a mortgage loan equation, that housing prices do not adjust to disequilibria in the market for housing loans. Hence, in a long run perspective the causation does not run from mortgage lending to housing prices. In the short run they find evidence of a contemporaneous bi-directional dependence.

Gerlach and Peng (2005) examine the interaction between credit to the private sector and residential property prices with a sample of quarterly data for Hong Kong from 1984 to 2001. Adopting a vector error correction framework they also find that the direction of causation goes from housing prices to private sector debt both in the long run and in the short run.

Fitzpatrick and McQuinn (2007) look at the interaction between housing prices

[2] See also Goodhart and Hofmann (2007).

and mortgage credit in Ireland between 1981 and 1999. They show that the two variables are mutually dependent in the long run, as well as in the short run.[3] Like Hofmann (2003), Fitzpatrick and McQuinn (2007) analyze cointegration within a single equation framework by adopting the Engle-Granger approach[4] (Engle and Granger, 1987).

A uni-equational approach is used to specify the error correction models for housing prices and debt before estimating them jointly by three stage least squares (3SLS).[5] The equations in the system are estimated jointly after each equation was specified separately. This modeling strategy is practical, but – as pointed out by Hammersland and Jacobsen (2008) – an equation system that is estimated simultaneously after the equations in the system have been reduced and specified separately may be inappropriate. This is because the single equation specifications will themselves be affected by the reduction process if we believe the variables in the system are jointly determined in the first place.

The results of Fitzpatrick and McQuinn (2007) is supported by Berlinghieri (2010) employing US quarterly data for the period 1977 to 2005. A two-step Engle-Granger approach is adopted and the interaction is found to go in both directions both in the long and the short run.

Oikarinen (2009) uses quarterly data for Finland (the sample is 1975 to 2006) to explore the mutual dependence between housing prices and borrowing. A Johansen analysis supports the existence of only one cointegrating vector and the direction of causation is found to be from household borrowing to housing prices in the long run. An impulse response analysis shows that the interaction between housing prices and credit grew to become important only after 1987, *i.e.* after the Finnish credit market was considered fully deregulated.

Making use of quarterly data for the period 1984-2009, Gimeno and Martínez-Carrascal (2010) study the interaction between housing prices and household borrowing in Spain. A multivariate cointegration analysis shows that the two variables are interdependent in the long run, *i.e.* that housing prices affect mortgage credit in the long run, and vice versa. Further, the loading factors imply that disequilibrium in the credit market leads to adjustments in both markets, while only housing prices error correct to disequilibrium constellations in the housing market. They do not report results on the short run dynamics, and they are therefore not able to study the dynamics of the system.

Our contribution is along the lines of Gimeno and Martínez-Carrascal (2010), but in addition to study the long run interaction, we model the dynamic interaction between the two markets, which is important for both policy evaluation and forecasts. Moreover, our study is first of its kind on Norwegian data: Existing models of Norwegian housing prices do not take explicit account of the feedback between

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[3] They do however not find a contemporaneous effect from housing prices on credit in the short run.

[4] Hofmann (2003) also considers Johansen analysis, but it is the results from the single equation tests that are retained for the short run specification.

[5] In addition to an equation for housing prices and one for household debt, Fitzpatrick and McQuinn (2007) adds an additional equation for the supply side of the housing market to their system. This equation is taken from a former study (McQuinn, 2004) and is not directly integrated in their full analysis.

the housing and credit markets.

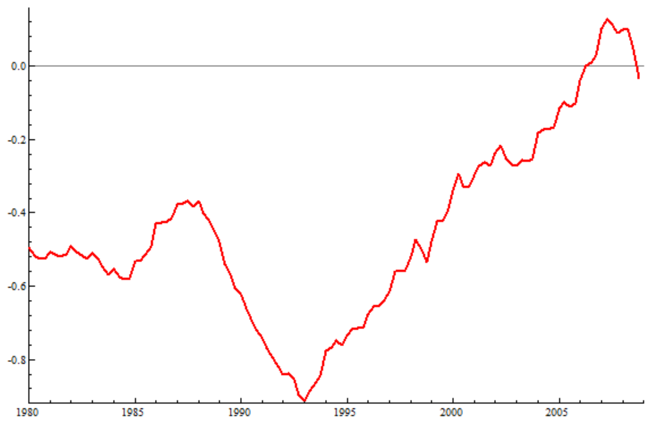
3 The Norwegian housing and credit markets

Since World War II Norway has seen two examples of a collapse of property prices being followed by imbalances in the real economy: The banking crisis that took place between 1988-1993 and the recent financial crisis.

Krogh (2010) gives a detailed account of the changes in the Norwegian credit market regulations and other major events in the period 1970-2008. This time span entails a period with strict credit market regulations in the 1970s, a gradual deregulation of these markets in the 1980s, followed by the banking crisis, and the subsequent development up to the advent of the current financial crisis.[6]

For our purpose it is important to note that also the housing market was deregulated in 1982, shortly before the credit market regulations were lifted. The combined effect of these liberalization processes was a boom in the real estate market, made possible and financed by a credit expansion. The problems facing the banking sector when the bubble burst became immense (Vale, 2004). After the end of the Norwegian banking crisis in 1993, real housing prices have grown almost consecutively until the financial meltdown of the previous decade (see Figure 1). Growing housing prices have been accompanied by a substantial expansion in real household debt (see Figure 2).

Figure 1: Log of real housing prices, 1980-2008

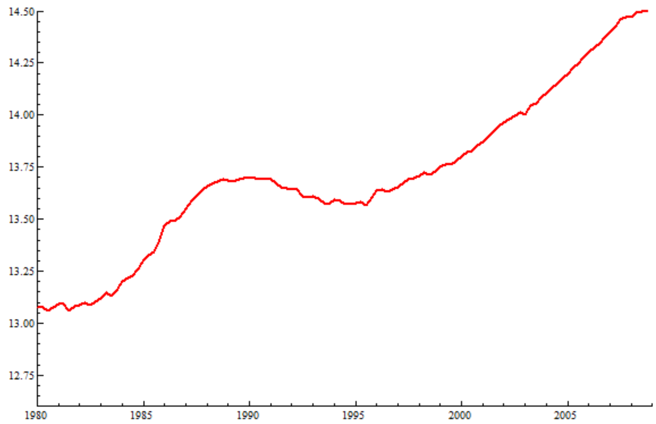


The historical episodes referred to above suggest there is an interdependency between the evolution of real housing prices and that of real household debt. For an impression of how housing price developments relate to the general macroeconomic picture in Norway, Figure 3 plots the four quarter growth in real housing prices against percentage deviations of GDP mainland Norway from trend.[7] A close link between economic activity and housing prices is apparent over the entire period, with

[6] See also the discussion in Jansen and Krogh (2011).

[7] GDP mainland Norway measures total production in Norway excluding extraction of oil and gas as well as other production related to this. The trend is constructed using an HP-filter with $\lambda = 40000$ (Source: Statistics Norway).

Figure 2: Log of real household debt, 1980-2008



a less pronounced correlation pattern the last few years. Goodhart and Hofmann (2007) argue that there is a tendency of changes in housing price growth to lead *peaks* and *troughs* in economic activity. This may suggest that turning points in the housing market are indicators of future economic developments. Figure 3 shows such a tendency for the case of Norway in the period after deregulation. Housing prices may affect economic activity through wealth effects on private consumption and a rise in house prices also raise the value of housing relative to construction costs, that is the Tobin q (Tobin, 1969) for residential investments. Another channel could be that housing prices contribute to the economic cycle by amplifying shocks in the credit market. It is evident from Figure 4, where we have plotted the four quarter growth in real housing prices against four quarter growth in real household borrowing, that the two series are closely linked and seem to move quite closely together.

Figure 3: GDP gap (left scale) and four quarter growth in real housing prices (right scale), 1985-2008

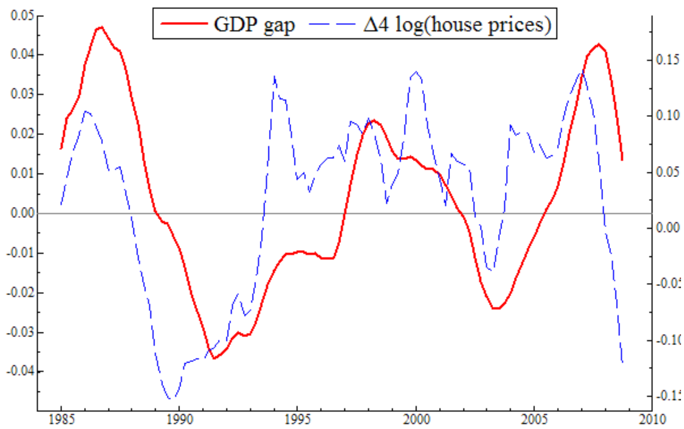
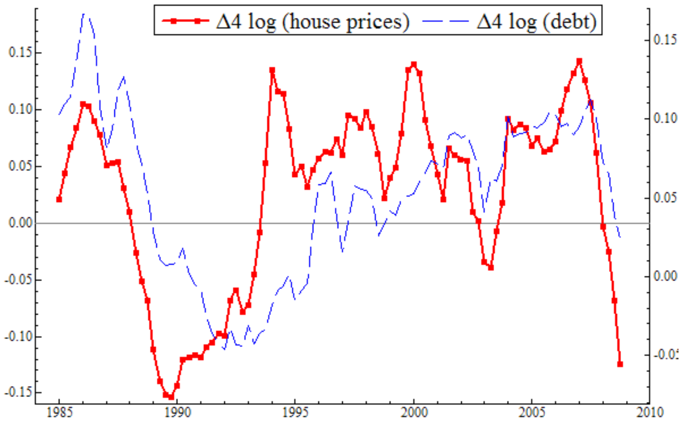


Figure 4: Four quarter growth in real housing prices (left scale) and in real household debt (right scale), 1985-2008



Previous studies of the credit and housing markets in Norway do not take the potential simultaneity between these two variables into account. For example, Jacobsen and Naug (2004, 2005) are two separate studies on household debt and housing prices, respectively. In Jacobsen and Naug (2005), household borrowing is not part of the cointegrated vector explaining housing prices.[8] It is, however, documented that the interest rate is an important determinant of housing prices. Jacobsen and Naug (2004) find that the interest rate is one of the fundamental factors explaining household borrowing. Since credit effects are not modeled explicitly in Jacobsen and Naug (2005), it may very well be the reason why the interest rate enters significantly and exerts a great impact on housing prices. In such a case, the interest rate is likely to be an indirect way of capturing credit effects and the coefficient of the interest rate is a gross effect.[9]

4 Long run analysis

While some studies have favored a multivariate approach to cointegration in a housing market-credit market set-up, few attempts have been made to unify this approach with a full study of the dynamic model that links these markets together. In the current paper we use a dynamic econometric model to assess the joint relationship between housing prices and credit, and in doing so we seek to answer one key question: Is there empirical evidence for the existence of a financial accelerator in the Norwegian housing market?

To investigate this issue we formulate a vector error correction model (VECM). This approach provides an opportunity to study long run determinants and short run dynamics in a unified framework. As the above discussion indicates, it seems

[8] Jacobsen and Naug (2005) tested for the significance of a credit variable in their specification, but found no significant effects.

[9] The household debt and housing price equations in SMM, which is a macroeconomic model at Norges Bank, are derived from those in Jacobsen and Naug (2004, 2005). The housing price equation of SMM has been modified to include a long run effect of credit, as mentioned in Andersen et al. (2008) and Andersen (2011). There is, however, as yet no further documentation available for the new equation.

plausible that the direction of causality is bi-directional. Resorting to a multivariate cointegration approach allow us to treat both variables as endogenous in the long run.

As a theoretical backdrop for the empirical modeling, we formulate the determination of real housing prices and real household debt at the aggregate level in a static long run equilibrium:

$$PH = f(H, YH, D) \tag{1}$$

where $\frac{\partial f}{\partial H} < 0, \frac{\partial f}{\partial YH} > 0, \frac{\partial f}{\partial D} > 0$ and

$$D = g(H, YH, R, PH, TH) \tag{2}$$

with $\frac{\partial g}{\partial H} > 0, \frac{\partial g}{\partial YH} > 0, \frac{\partial g}{\partial R} < 0, \frac{\partial g}{\partial PH} > 0, \frac{\partial g}{\partial TH} > 0$. In (1) and (2) PH denotes real housing prices, YH represents real disposable income for the household sector (excluding dividends), R is the real interest rate after tax, D is real household debt, H is the housing stock and TH is housing turnover.

Equation (1) is an inverted demand equation for housing services, and expresses market clearing prices for any given housing stock. This is methodologically in line with for example the seminal contribution by Hendry (1984). Equation (1) describes housing prices as an increasing function of disposable income and household debt, while a greater supply of housing services is expected to push housing prices down.

Equation (2) is an extended version of Fitzpatrick and McQuinn (2007) . It defines household debt as a function of the housing stock, housing prices, the interest rate, disposable income and the housing turnover. In our specification, the housing stock and the housing turnover are additional explanatory variables.

The equation system defined by (1) and (2) assumes that there are no direct causal link that tie housing prices to the real interest rate in the long run. One might question the exclusion of direct interest rate effects in Equation (1). However, as almost all mortgage debt in Norway are loans with flexible interest rates, a change in interest rates will immediately feed into the disposable income for households, and it is likely to pick up the main effect of interest rates on demand for housing. In addition, interest rate effects are preserved by the presence of the credit aggregate which captures the effect on housing prices from a change in the cost of financing.[10] A similar assumption is made by Gimeno and Martínez-Carrascal (2010) and Fitzpatrick and McQuinn (2007).

In the empirical analysis, a semi logarithmic transformation of the system defined by (1) and (2) is considered. All data are seasonally unadjusted and in what follows, small letters indicate that the variables are measured on a logarithmic scale.[11] All monetary variables are measured in real terms, having been deflated by the consumption deflator. Our sample covers the period 1986q2-2008q4. The reason for not starting earlier is that we only have data for the number of house

[10] Recall that our interpretation of Jacobsen and Naug (2004, 2005), that the interest rate effect on housing prices is likely to be an indirect way of capturing credit effects on that variable, *i.e.* the coefficient of the interest rate is a gross effect due to the omitted variable.

[11] For a detailed data description, see Appendix A. The log transformation is applied to all variables in (1) and (2), except the real after tax interest rate.

transactions from 1985q1 and also because the housing price data are less reliable in the period prior to this. In addition the deregulation of the two markets that took place in the early 1980's is likely to have altered the functioning of both, so that a different model would probably be more appropriate if we want to consider the period prior to the deregulation.

An important issue when working with time series data is whether the variables are stationary or not. Granger and Newbold (1974) showed how regressions on non-stationary variables might lead to *non-sense* or *spurious* regressions because the residuals from such a regression will in general be non-stationary unless a linear combination of the variables are stationary. To investigate the order of integration of the data, standard Augmented Dickey-Fuller (ADF) tests are performed. In Table B.1 in Appendix B the results from the ADF-tests are displayed. As a guidance for choosing the optimal lag truncation we have relied on Akaike's information criterion (AIC).[12] It is evident that most of the time series in our sample are integrated of order one. The real after tax interest rate appears to be stationary, which does not impose any problems for the empirical analysis. The tests indicate that real housing prices are integrated of order two, which also seems to be the case for the housing stock. As far as the credit aggregate is concerned, the results are more ambiguous.[13]

The fact that the tests indicate that some of the series are integrated of an order greater than one, could potentially cause some problems of misinterpretation when we carry out the analysis using cointegration techniques for the I(1) case. Still, we have chosen to do so. There is however supporting evidence for this approach in that we find – as we report below – that the residuals in the final empirical model as well as the error correction terms of that model turn out to be stationary.[14]

In the cointegration analysis, we start with a fifth order VAR in real housing prices, real household debt and real disposable income, conditioning on the real after tax interest rate, the housing turnover and the housing stock. Re-expressed in an error correction representation, we get the following expression:

$$\Delta \mathbf{x}_t = \mathbf{\Pi} \mathbf{y}_{t-1} + \sum_{i=1}^{4} \mathbf{\Gamma}_i \Delta \mathbf{x}_{t-i} + \sum_{i=0}^{4} \boldsymbol{\zeta}_i \Delta \mathbf{z}_{t-i} + \boldsymbol{\delta} t + \boldsymbol{\epsilon}_t \tag{3}$$

where $\boldsymbol{\epsilon} \sim \mathrm{N}(0,\mathbf{\Sigma})$, $\mathbf{x}$ is a 3×1 vector comprising the endogenous variables ph, d and yh. $\mathbf{y} = (\mathbf{x}', \mathbf{z}')'$ is a $(3+3) \times 1$ vector where $\mathbf{z}$ is a 3×1 vector composed of the exogenous variables R, th and h. t is a linear trend, and $\mathbf{\Pi}$, $\mathbf{\Gamma}_i$, and $\boldsymbol{\zeta}_i$ are the corresponding coefficient matrices and $\boldsymbol{\delta}$ is a vector of (trend) coefficients. The trend is restricted to lie in the cointegrating space.

[12] For all time series to be tested, we initially started with eight lags in the first differences. While the start of the sample is 1985q1 for the ADF-test, it is 1986q2 in the later analysis. The reason for this is that we include five lags in the VAR when testing for cointegration, so that the earliest observation used to construct the lags is 1985q1.

[13] The test rejects non-stationarity in levels. However, when testing the first and second differences it is shown that Δd is I(1), which would imply that d is in fact an I(2) variable. It should be noted that the ADF tests have low power in this case.

[14] The test supports stationarity of the error correction term for housing prices at the 10 percent level. For the error correction term of household borrowing the test indicates stationarity at the 5 percent level.

Table 1: Trace test for cointegration [a]

$Eigenvalue: \lambda_i$	H_0	H_A	λ_{trace}	5%-critical value[b]
0.38909	$r = 0$	$r \geq 1$	86.587	64.48
0.22269	$r \leq 1$	$r \geq 2$	41.472	40.95
0.18681	$r \leq 2$	$r \geq 3$	18.818	20.89

Diagnostics[c]	Test statistic	Value[p-value]
Vector AR 1-5 test:	F(45,146)	1.10585 [0.3902]
Vector Normality test:	$\chi^2(6)$	7.7765 [0.2549]
Vector Hetero test:	F(228,139)	0.53408 [1.0000]
Estimation period:	1986q2-2008q4	

[a] Endogenous variables: Real housing prices, real household debt and real disposable income. Restricted variables: Real interest rate after tax, housing turnover, housing stock and a trend. Unrestricted variables: Constant and centered seasonal dummies for the first three quarters.
[b] Critical values are obtained from Table 13 in Doornik (2003) - with 3 exogenous variables.
[c] See Doornik and Hendry (2009).

The modeling assumption that the variables are at most integrated of order 1 (*i.e.* I(0) or I(1)) and the hypothesis of joint causation imply that the multivariate cointegration approach of Johansen (1988) is the appropriate estimation procedure. This system based method allows several variables to be treated as endogenous in the cointegration space, and for that reason provides a suitable framework for formally testing the hypothesis that housing prices and household debt are mutually dependent in the long run.

According to AIC the VAR-model should include 5 lags in the endogenous variables (tests are not reported here). To determine the lag length of the exogenous variables included in the information set, we report a series of Wald F-tests (see the lower part of Table B.3 in Appendix B). Judged by these tests it is sufficient to include only one lag in each of the exogenous variables. The same lag structure is implied if we instead rely on AIC.

A trace test for the order of cointegration (Johansen, 1988) is conducted to determine the number of long run relationships that can be derived between the variables in our model. The trace test determines the rank of a matrix $\mathbf{\Pi}^e = (\mathbf{\Pi}, \boldsymbol{\delta})$ and gives the number of independent linear combinations between the variables that are stationary. Table 1 displays the results. When correcting for the inclusion of exogenous variables (see Harbo et al. (1998) and Doornik (2003)), the test indicates the existence of two cointegrating vectors. The model is well specified – diagnostics imply that the residuals are neither heteroskedastic nor autocorrelated, and normality is not rejected.

We follow Johansen (1988) and define $\mathbf{\Pi}^e = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\beta}$ is a $(p+q+1) \times r$ matrix and $\boldsymbol{\alpha}$ is a $p \times r$ matrix corresponding to the long run coefficients and loading factors respectively. The rank of the $\mathbf{\Pi}^e$ matrix is denoted by r, while p refers to number of endogenous variables and $q+1$ is the number of exogenous variables (including the deterministic trend, which is restricted to lie in the cointegration space). Given that the rank of $\mathbf{\Pi}^e$ is two, with three endogenous and three exogenous variables, the cointegrating part of equation (3) takes on the following form:

(4)

$$\boldsymbol{\alpha\beta}'\mathbf{y} = \begin{pmatrix} \alpha_{1,ph} & \alpha_{1,d} \\ \alpha_{2,ph} & \alpha_{2,d} \\ \alpha_{3,ph} & \alpha_{3,d} \end{pmatrix} \begin{pmatrix} \beta_{ph,1} & \beta_{d,1} & \beta_{yh,1} & \beta_{R,1} & \beta_{th,1} & \beta_{h,1} & \beta_{t,1} \\ \beta_{ph,2} & \beta_{d,2} & \beta_{yh,2} & \beta_{R,2} & \beta_{th,2} & \beta_{h,2} & \beta_{t,2} \end{pmatrix} \begin{pmatrix} ph \\ d \\ yh \\ R \\ th \\ h \\ t \end{pmatrix}$$

Identification can be achieved by imposing two restrictions in each vector. We start by normalizing on real housing prices in the first vector ($\beta_{ph,1} = 1$) and real household debt in the other ($\beta_{d,2} = 1$). In addition, it is assumed that the real interest rate after tax has no direct effect on real housing prices ($\beta_{R,1} = 0$).[15] As mentioned above, this does not imply that a change in the interest rate will not affect housing prices, but we assume that interest rate effects are captured by changes in disposable income and through the credit channel.

The final restriction we impose is based on the earlier argument that it is the value of the housing capital – and not simply housing prices – which determines the size of the collateral. To incorporate this into the empirical framework, we assume that a change in either the housing stock or housing prices have the same effect on household debt ($\beta_{ph,2} = \beta_{h,2}$).

Having identified the two cointegrated vectors as a housing price and credit equation, we carry out tests of overidentifying restrictions. The results from the remaining steps are documented in Table 2 and Table 3 below.[16] In Panel 1 the trend variable is dropped from both equations ($\beta_{t,1} = \beta_{t,2} = 0$) which correspond to two testable overidentifying restrictions. Next, in Panel 2, we are omitting the housing turnover variable from the vector associated with the real housing price ($\beta_{th,1} = 0$). In Panel 3 there is no effect of disequilibrium in the housing market on household debt ($\alpha_{2,ph} = 0$), whereas Panel 4 shows the case with no effect of real disposable income on household debt ($\beta_{yh,2} = 0$). The final Panel 5 shows the result when we impose that the loadings of both cointegrating vectors with respect to income are zero ($\alpha_{3,ph} = \alpha_{3,d} = 0$), *i.e.* weak exogeneity of income with respect to the long run coefficients. All the individual restrictions are supported by the data and the p-value for the joint test of all restrictions is 0.3.

The coefficients reported in Panel 5 in Table 2 describe the two long run relationships we find for housing prices and household debt, respectively. The loadings or error correction terms are reported in Table 3. Our results support the hypothesis that housing prices and household borrowing are mutually dependent in the long run. All long-run coefficients have the expected signs and are significant at conventional significance levels.[17] In Table B.2 in Appendix B, we report ADF tests

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[15] If we alternatively obtain exact identification of the first cointegrating equation by excluding the turnover, *i.e.* $\beta_{th,1} = 0$, the coefficient of the interest rate comes out highly insignificant. Hence, excluding the interest rate at that stage is accepted by the data and this leaves us with the same vector as those obtained i Panel 2 of Table 2.

[16] Standard errors are reported in parantheses below the estimated coefficients.

[17] The interest rate is the only exception. However, using a one sided test, which appears to be

Table 2: Testing steady-state hypotheses.

The just identified house price and debt equations is defined by

$\beta_{ph,1} = 1$, $\beta_{d,2} = 1$, $\beta_{R,1} = 0$, $\beta_{ph,2} = \beta_{h,2}$ in (4)

$LogL = 844.634$

Panel 1: Testing no trend ($\beta_{t,1} = \beta_{t,2} = 0$)

$$ph = \underset{(0.11)}{0.82}d + \underset{(0.34)}{1.62}yh + \underset{(0.10)}{0.04}th - \underset{(0.57)}{2.54}h$$

$$d = 1.52ph - \underset{(0.37)}{1.43}yh - \underset{(0.58)}{0.78}R + \underset{(0.16)}{0.08}th + \underset{(0.10)}{1.52}h$$

$LogL = 842.845$, $\chi^2(2) = 3.81[0.15]$

Panel 2: No effect of housing turnover on house prices ($\beta_{th,1} = 0$)

$$ph = \underset{(0.08)}{0.77}d + \underset{(0.22)}{1.43}\,yh - \underset{(0.40)}{2.07}h$$

$$d = 1.54ph - \underset{(0.18)}{1.48}yh - \underset{(0.40)}{0.54}R + \underset{(0.05)}{0.10}th + \underset{(0.07)}{1.54}h$$

$LogL = 842.834$, $\chi^2(1) = 0.02[0.88]$, $\chi^2(3) = 3.84[0.28]$

Panel 3: No effect of disequilibrium housing prices on household debt($\alpha_{2,ph} = 0$)

$$ph = \underset{(0.19)}{0.84}d + \underset{(0.65)}{1.67}yh - \underset{(1.18)}{2.58}h$$

$$d = 1.08ph - \underset{(0.85)}{1.18}yh - \underset{(2.35)}{3.98}R + \underset{(0.28)}{0.56}th + \underset{(0.30)}{1.08}h$$

$LogL = 842.276$, $\chi^2(1) = 1.12[0.29]$, $\chi^2(4) = 4.95[0.29]$

Panel 4: No effect of real disposable income on household debt ($\beta_{yh,2} = 0$)

$$ph = \underset{(0.19)}{0.86}d + \underset{(0.64)}{1.42}yh - \underset{(1.16)}{2.33}h$$

$$d = 0.78ph - \underset{(1.87)}{2.81}R + \underset{(0.15)}{0.24}th + \underset{(0.15)}{0.78}h$$

LogL = 841.323 , $\chi^2(1) = 1.12[0.29]$, $\chi^2(5) = 6.86[0.23]$

Panel 5: Imposing weak exogeneity of income with respect to the long run coefficients ($\alpha_{3,ph} = \alpha_{3,d} = 0$) :

$$ph = \underset{(0.19)}{0.98}d + \underset{(0.63)}{1.69}yh - \underset{(1.15)}{3.03}h$$

$$d = 0.76ph - \underset{(1.79)}{2.74}R + \underset{(0.15)}{0.28}th + \underset{(0.16)}{0.76}h$$

$LogL = 840.529$, $\chi^2(2) = 1.59[0.451]$, $\chi^2(7) = 8.44[0.30]$

The sample is 1986q2 to 2008q4, 91 observations.

for the residuals from the estimated long run relationships of housing prices and credit respectively. The tests imply that the residuals from the credit equation are stationary. We find that the hypothesis of non-stationarity of the residuals of the credit equation can be rejected at the 5% level whereas the housing price equation is rejected at the 10% level. We interpret this as support for stationarity of both equations which corroborates their interpretation as cointegrating equations. The result for the housing price equation is in part a consequence of the large deviations from equilibrium we observe in the period prior to, during and after the banking crisis of the 1990's. This point is also stressed by Koetter and Poghosyan (2010), Goodhart and Hofmann (2007) and Zhu (2005) who emphasize that housing prices are more likely to have persistent deviations from their long run equilibrium than

meaningful, it is found to be significant at the 10 % level (p-value = 0.068). The fact that it is also highly significant from an economic point of view suggests that it should not be excluded.

Table 3: Loading factors.

Loading	Panel 1	Panel 2	Panel 3	Panel 4	Panel 5
$\alpha_{1,ph}$	-0.54 0.14	-0.76 0.20	-0.22 0.04	-0.22 0.04	-0.24 0.04
$\alpha_{1,d}$	-0.25 0.09	-0.40 0.13	-0.06 0.02	-0.08 0.03	-0.10 0.03
$\alpha_{2,ph}$	-0.08 0.07	-0.13 0.07	0 −−	0 −−	0 −−
$\alpha_{2,d}$	-0.08 0.05	-0.10 0.07	-0.04 0.01	-0.05 0.01	-0.04 0.01
$\alpha_{3,ph}$	0.28 0.10	0.40 0.14	-0.04 0.03	-0.05 0.03	0 −−
$\alpha_{3,d}$	0.21 0.06	0.29 0.06	0.002 0.02	0.01 0.02	0 −−

other asset prices.

The semi-elasticity of household borrowing with respect to the real interest rate after tax is -2.74, implying that a one percentage point increase in the real interest rate will decrease household borrowing by almost three percent in the long run. This is lower (in absolute value) than the estimate found for Spain by Gimeno and Martínez-Carrascal (2010) [-9.51] who consider nominal instead of real interest rates. It is however greater than the estimates found by Brissimis and Vlassopoulos (2009) [-0.039] for Greece and Fitzpatrick and McQuinn (2007) [0.007] for Ireland who both look at real interest rates. Even though there is no direct causal link between real housing prices and the real interest rate in our model, a higher interest rate implies that housing prices will fall as it reduces the demand for housing by altering the credit variable, which is found to be highly significant in the housing price equation.

The estimated elasticity of housing prices with respect to household debt is 0.98. This is lower than the elasticity reported by Fitzpatrick and McQuinn (2007), but higher than the estimate in Gimeno and Martínez-Carrascal (2010). We find that the credit aggregate exercises a greater impact on housing prices than do housing prices on credit in a long run perspective, a result that parallels the finding of Fitzpatrick and McQuinn (2007). A one percent increase in housing prices will increase household borrowing by 0.76% in the long run.

The loadings imply that both housing prices and household debt error correct when the latter departs from the value implied by its fundamentals. Moreover, the analysis indicates that only housing prices error correct when deviating from its steady state level. This result is supported by Gimeno and Martínez-Carrascal (2010).

5 Short run dynamics

In this section we exploit the results from the cointegration analysis, and embed the residuals from Section 4 as error correction terms in the short run model. At the outset we start with a general model formulation consistent with the VAR analysis in Section 4:

$$\Delta ph_t = \beta_0 + \sum_{i=1}^{4} \beta_{1i}\Delta ph_{t-i} + \sum_{i=0}^{4} \beta_{2i}\Delta d_{t-i} + \sum_{i=0}^{4} \boldsymbol{\eta}_i \Delta \mathbf{z}^*_{t-i} + \alpha_{1,ph} ECM^{ph}_{t-1} + \alpha_{1,d} ECM^{d}_{t-1}$$
$$+ CSeasonal_t + CSeasonal_{t-1} + CSeasonal_{t-2} + \epsilon_{ph,t} \quad (5)$$

$$\Delta d_t = \gamma_0 + \sum_{i=1}^{4} \gamma_{1i}\Delta d_{t-i} + \sum_{i=0}^{4} \gamma_{2i}\Delta ph_{t-i} + \sum_{i=0}^{4} \boldsymbol{\psi}_i \Delta \mathbf{z}^*_{t-i} + \alpha_{2,ph} ECM^{ph}_{t-1} + \alpha_{2,d} ECM^{d}_{t-1}$$
$$+ CSeasonal_t + CSeasonal_{t-1} + CSeasonal_{t-2} + \epsilon_{d,t} \quad (6)$$

where $\mathbf{z}^*$ is a vector representation of the exogenous variables in the system, and β_{ji}, γ_{ji}, $\boldsymbol{\eta}_i$ and $\boldsymbol{\psi}_i$ are the short run coefficients. Since the housing stock adjusts slowly, it is assumed to be fixed in the short run and is not part of the vector $\mathbf{z}^*$. However, we supplement the short run dynamics by including the variable E, which measures households expectations about future developments in their personal economy and the macroeconomy. Hence, $\mathbf{z}^*= (R, th, E, yh)$. The expectations variable is only collected from 1992q3 and is set to 0 in the period prior to this. The expectations variable has previously been adopted by Jacobsen and Naug (2005). They find a positive and significant short run effect of expectations on housing prices in a single equation framework.

The system is estimated by FIML, where the errors are assumed to be normally distributed with zero expectation and variance-covariance matrix $\boldsymbol{\Omega}$. To identify the two equations in the system, the real after-tax interest rate is excluded at all lags from the equation for real housing prices and the real household debt equation is identified by omitting the contemporaneous value of the expectations variable. We have also left out the contemporaneous value of the real after-tax interest rate in the credit equation, as it is likely that it takes some time from the interest rate is changed until it has an effect on credit. Also, this produced a model that was more interpretable from a theoretical perspective.[18]

A parsimonious model is found by stepwise elimination of insignificant variables in the system. In the preferred model, we have chosen to retain some variables ,which are relevant from a priori theoretical considerations, although they should have been excluded at the early stages of the reduction process had we followed a strict general to specific procedure. By doing this, we have arrived at a more theoretically and intuitively appealing model formulation than we would have obtained otherwise.[19]

Unlike previous studies (c.f. Fitzpatrick and McQuinn (2007) and Brissimis and Vlassopoulos (2009)), the top-down approach applied in this paper consists of modeling the system simultaneously at all steps in the reduction process. Even though some weight has been attached to theoretical priors in the reduction process, only insignificant variables have been eliminated at each stage. At every step we have made sure that the Gaussian properties of the residuals were retained according

[18] Including the contemporaneous interest rate in the specification made the effect from this variable positive.

[19] Results from adopting a more orthodox GETS (general-to-specific) strategy, see Hoover and Perez (1999) and Hendry and Krolzig (1999, 2001), are available upon request,

to the diagnostic tests and that the imposed restrictions were supported by the data. The results are presented in Table 4.

Table 4: Short run dynamics [a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.542	7.71	0.048	6.39
Δd_t	0.859	2.25	-	-
Δd_{t-1}	-	-	0.173	1.88
Δd_{t-3}	0.309	2.32	-	-
Δph_{t-4}	0.389	4.88	-	-
Δyh_{t-3}	-	-	0.197	3.31
ΔE_t	0.093	4.40	-	-
ΔE_{t-1}	0.098	4.41	-	-
ΔE_{t-2}	0.055	2.40	-	-
ΔR_{t-4}	-	-	-0.258	2.16
ECM^{ph}_{t-1}	-0.175	7.82	-	-
ECM^{d}_{t-1}	-0.059	2.23	-0.046	6.11
Dummy, q1	0.022	3.75	-0.004	1.18
Dummy, q2	0.021	3.65	-0.00001	0.02
Dummy, q3	0.012	2.05	-0.007	2.05
Sargan	$\chi^2(48) =$	61.011 [0.0984]		
Log likelihood	560.26			
σ	0.0143		0.0098	
Diagnostics[b]	Test statistic	Value [p-value]		
Vector EGE-AR 1-5 test:	F(20,140)	0.902 [0.5852]		
Vector Normality test:	$\chi^2(4)$	5.3371 [0.2544]		
Vector hetero test:	F(180,54)	0.5864 [0.9950]		
Estimation Method	FIML			
Sample	1986q2-2008q4			

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009).

An inspection of Table 4 reveales that liquidity effects are important for housing price fluctuations also in the short run. We do not find any direct short run effect running from household debt to housing prices. It is, however, clear that the credit aggregate will be influenced by housing prices through the error correction term present in the credit equation. Consistent with the cointegration analysis, the short run analysis indicates that both housing prices and household debt error correct when household debt is high relative to its stable long run equilibrium and that only housing prices error correct when departing from their fundamentals. Our results suggest that if housing prices depart from their long run equilibrium by 1%, housing prices will fall by -0.175%. This is greater than what is found by Jacobsen and Naug (2005)[20] $[-0.12\%]$, but lower than the estimate reported by Fitzpatrick

[20] Jacobsen and Naug (2005) only consider housing prices and not the interaction between housing prices and household debt.

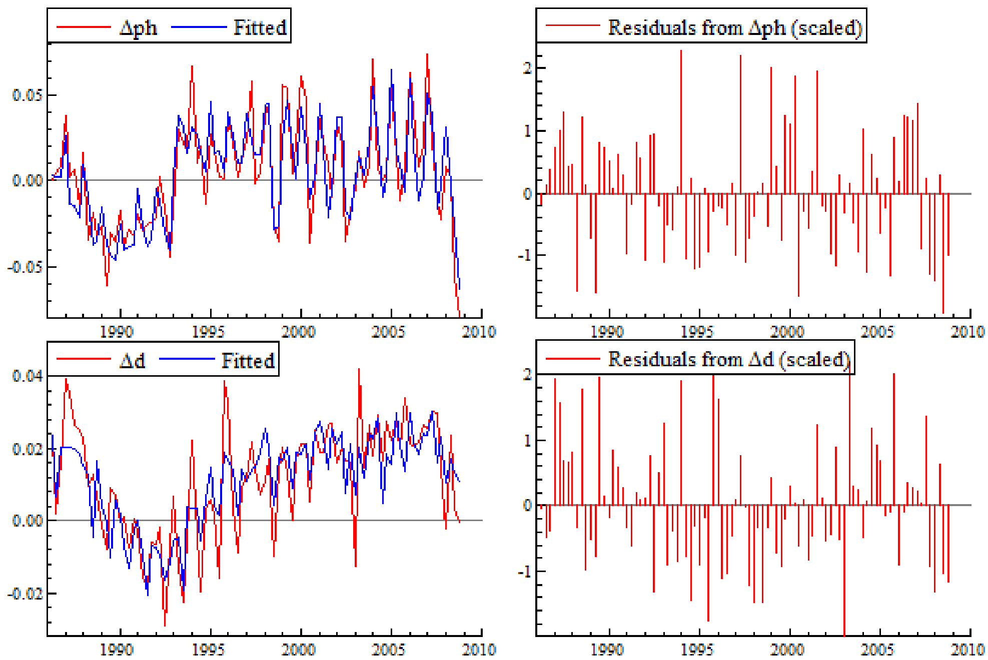
and McQuinn (2007) [−0.25%].

Like Jacobsen and Naug (2004, 2005) and Fitzpatrick and McQuinn (2007), we find that the credit aggregate has a slower adjustment towards equilibrium when departing from its fundamentals than housing prices. This is not a very surprising finding, in light of the fact that the volume of debt is not easily changed over night. Gimeno and Martínez-Carrascal (2010), however, find the opposite to be the case for Spain.

All estimated coefficients have the expected signs, and interestingly we document a rather great impact on housing prices of changes in expectations. The full effect is reached after three quarters, *i.e.*, when there has been a change of 'mood'. As anticipated, our estimation results show that the interest rate has a negative impact on household borrowing (and therefore indirectly on housing prices) and the income variable lagged three quarters enters the credit equation significantly with an expected positive sign. The diagnostics indicate that the model is well specified and we find support for the imposed restrictions (p-value = 0.0984). The residuals from the two estimated equations are clearly stationary (see Table B.2 in Appendix B).

In Figure 5 fitted values and residuals from the two equations are shown. The evolution of housing prices implied by our model closely follows the actual development. The debt equation also has a good fit, but deviates more from the actual development than do the aforementioned equation for housing prices. In the next section, we discuss the dynamic responses when the system is hit by different exogenous shocks.

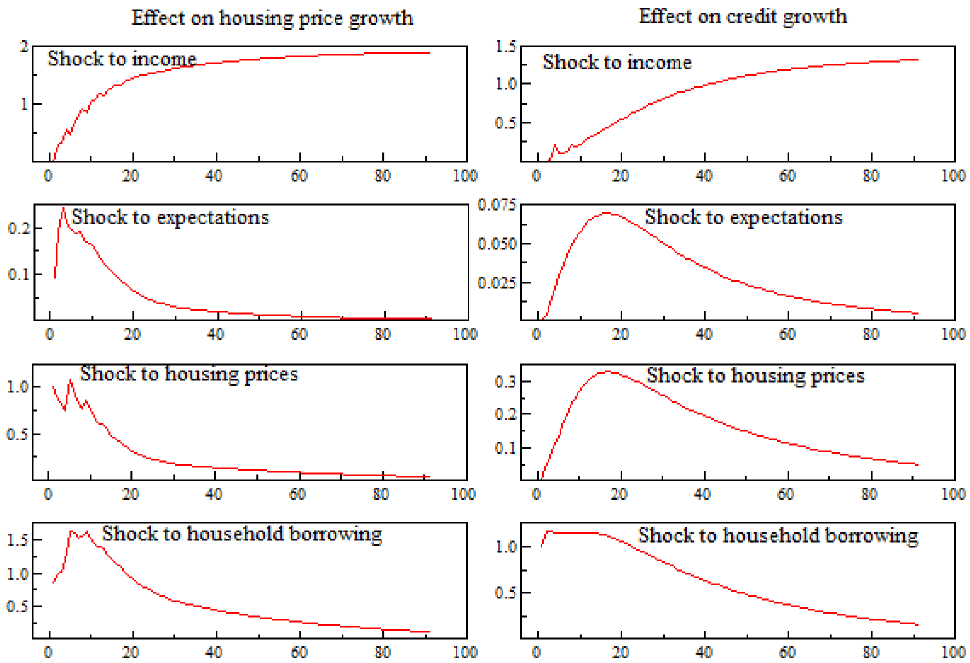
Figure 5: Fitted values and residuals, 1986q2-2008q4



6 Impulse response analysis

In this section we look at impulse responses to see how shocks are transmitted and possibly amplified in our dynamic model.[21] Figure 6 illustrates the cumulative multipliers when we shock the variables in the system. The figure demonstrates how shocks in one of the markets are propagated and amplified through an endogenous feedback mechanism.

Figure 6: Permanent shocks



A one percentage point increase in income growth will affect neither the growth in housing prices nor household debt in the period of the shock. After one quarter, however, housing prices start adjusting towards its new equilibrium (remember that a higher income increases the fundamental value of housing prices). In the next period this spills over to the credit market through the error correction term. As the cumulative multipliers illustrate, both the growth in housing prices and credit continue to grow before the growth eventually slows down and stabilizes. A permanent increase in income, which is one of the long run determinants of housing prices, will change housing prices and credit period after period as they keep adjusting towards their continuously changing long run equilibrium. The dynamic process, illustrated in Figure 6, clearly indicates that the relationship between housing prices and credit is mutually self-reinforcing. First, a higher income leads to increased property valuations, which raises the value of the collateral. This spills over to an increased credit supply, stimulating housing prices further and so on.

Other authors have emphasized the importance of expectations about future economic prospects for the mutual relationship between housing prices and credit,

[21] We introduce a separate equation for the variable to be shocked, so that we can give the error in that equation an exogenous impulse. A shock in one of the residuals might however also affect residuals in other equations if they are correlated. However, we find that the empirical variance-covariance matrix is nearly diagonal which validates our approach.

see *e.g.* Goodhart and Hofmann (2007).[22] Our analysis support this hypothesis by formally incorporating an explicit measure of households expectations into the analysis and showing that a shock to these expectations might generate a housing price-credit spiral. To be precise, the estimated model implies that an increase in the consumer confidence indicator by one index point (as households become more optimistic) will lead to an immediate increase in housing price growth by almost 0.1 percentage points, before reaching its peak after three quarters. At this point the cumulative increase in housing price growth amounts to approximately 0.25 percentage points. Also, the housing price growth has triggered a growth in real household debt. In a long run perspective the entire shock is absorbed, as a continuous change in growth rates can only be driven by a change in one of the fundamentals.

An exogenous shock in the credit aggregate will change housing price growth by 0.86 percentage point at the time of the shock. The change in housing prices result in a further change in the credit growth in the next period, as the collateral value has increased. This again stimulates further growth in housing prices and credit. This tendency seems to continue for about two years before the error correction term dominates and the effect of the shock gradually dissipates. In the long run there is of course no change in neither of the growth rates as none of the fundamentals have been changed. Shocking housing price growth yields qualitative effects that are similar to the above described effects, and will of course not change any of the variables in the long run.

7 Conclusions

In this paper we first show that the cointegration analysis supports two long run relationships: one for housing prices and one for household debt. Also, household income can be considered weakly exogeneous with respect to the long run parameters. We find that housing prices depend on household borrowing, real disposable income and the housing stock in the long run, whereas real household debt is driven by the value of housing capital (housing prices times the housing stock), the real interest rate and the housing turnover. Housing prices and household debt are mutually dependent as both appear in the long run equation for the other. This suggests that there are feedback effects between the two in the long run. That said, housing prices are equilibrium correcting to deviations from both long run equations, whereas household debt adjusts only to debt disequilibria.

Second, we embed the long run equations from the cointegration analysis in a simultaneous system explaining the changes in housing prices and debt, following a general to specific strategy. The equations are estimated simultaneously by full information maximum likelihood methods and insignificant variables are removed stepwise from the two equations. The estimation results suggest that the credit

[22] As they put it (Goodhart and Hofmann, 2007, p.15): "A rise in house prices, caused by more optimistic expectations about future economic prospects, raises the borrowing capacity of firms and households by increasing the value of collateral. Part of the additional available credit may also be used to purchase property, pushing up property prices even further, so that a self-reinforcing process may evolve".

aggregate is important for housing price dynamics, but that housing prices only affect household borrowing through the error correction term.

Third, a consumer confidence indicator measuring households' expectations concerning future developments in their private economy as well as the Norwegian macro economy is incorporated into our framework. This variable explicitly picks up expectations about future economic conditions and is shown to enter significantly in the housing price equation in the short run.

Finally, the dynamic multipliers of the impulse response analysis provides clear evidence for the existence of a credit-housing price spiral in Norway. Higher housing prices result in higher credit growth due to collateral effects, which again spurs housing price growth and so on, showing that there indeed is a financial accelerator at work.

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A Appendix A Data definitions

All data are seasonally unadjusted and measured on a quarterly basis. Except for the interest rate and the consumer confidence indicator all variables are transformed to log scale in the empirical analysis. Variable definitions and a brief description of the data are listed below.

pc: The consumption deflator in the National Accounts. Source: Statistics Norway.
ph: Hedonic housing price index measuring average housing prices in Norway. The index is calculated on the basis of data on sales in the second hand market. Statistics Norway officially started publishing housing price data in 1992. Prior to 1992 an unofficial index based on similar sources and compiled at Statistics Norway is used. The housing price index is deflated by pc. Source: Statistics Norway.
d: Total amount of outstanding gross household debt. Deflated by pc. Source: Statistics Norway.
yh: Households' disposable income, excluding equity income. Deflated by pc. Source: Statistics Norway.
h: Real housing stock measured in fixed prices. Measures the total stock of housing in Norway and is calculated according to the perpetual inventory method. Source: Statistics Norway.
th: The housing turnover measures the number of housing transactions. Source: Statistics Norway.
E: The expectations variable is taken from TNS Gallup and can be seen as a consumer confidence indicator. It is based on a survey, where average score can range between -100 and 100. In this paper we have normalized the variable to lie between -1 and 1. The indicator measures households expectations concerning the state of the economy and the development in their personal economy. Source: TNS-Gallup.
i: Nominal after-tax interest rate paid by households on loans in private financial institutions. Source: Statistics Norway.
p: Consumer Price Index. Source: Statistics Norway.
τ : Capital tax rate. After a tax reform in 1992 τ has been constant at 0.28. Source: Statistics Norway.
R: Real after-tax interest rate $(i * (1 - \tau) - \Delta_4 p)$.

B Tables

Table B.1: Augmented Dickey-Fueller test for order of integration[a]

Levels					
Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
ph	−3.117	−3.46	8	Yes	Yes
d[b]	−3.589	−3.46	8	Yes	No
h	−2.352	−3.46	5	Yes	No
yh	−2.156	−2.89	5	No	Yes
th[c]	−2.458	−3.46	6	Yes	Yes
E[d]	−1.558	−2.89	0	No	Yes
R[e]	−4.103	−3.46	0	Yes	No
First difference					
Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
Δph	−1.865	−2.89	6	No	Yes
Δd[f]	−2.426	−2.89	3	No	No
Δh	−2.605	−2.89	4	No	No
Δyh	−3.929	−2.89	4	No	Yes
Δth	−8.708	−2.89	1	No	Yes
ΔE	−6.097	−2.89	0	No	Yes
Second difference					
Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
Δ^2ph	−4.480	−2.89	5	No	Yes
Δ^2d	−12.69	−2.89	2	No	No
Δ^2h	−2.716	−2.89	3	No	No

[a] We only have data for the turnover from 1985q1 and 5 lags are included in the VAR. For that reason the estimations starts in 1986q2. To construct the real interest rate, we need four additional lags in CPI, so our data set covers the period 1984q1-2008q4. For this reason the variables are tested over the same period as the estimation is carried out, 1986q2-2008q4, since 9 lags (8 in first differences) are used as a starting point in the tests. For obvious reasons one and two additional observations are lost when testing the variables in first and second differences, respectively.

[b] If seasonal dummies are included, d non stationarity can not be rejected.

[c] Data are available only from 1985q1, so variable is tested over the sample 1987q2-2008q4.

[d] Data are available only from 1992q3, so variable is tested over the sample 1994q3-2008q4.

[e] Estimation starts in 1987q2 due to lags. The variable is also stationary if we instead start in 1986q2.

[f] At two lags the test shows stationarity of Δd, i.e. d is I(1). Same result applies when seasonal dummies are added, i.e. that d is I(1) when two lags are used instead of three.

Table B.2: Augmented Dickey-Fueller tests for residuals[a]

Levels					
Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
ecm_{ph}	−3.124	−3.46	6	Yes	No
ecm_d	−4.122	−3.46	8	Yes	No
$\varepsilon_{\Delta ph}$	−9.307	−2.89	0	No	No
$\varepsilon_{\Delta d}$	−7.945	−2.89	1	No	No

[a] We only have data for ecm_d from 1985q1 because it includes the turnover. For this reason the two error correction terms are tested over the sample 1987q3-2008q4. The residuals from the short run system is tested over the period 1988q3-2008q4 since we only obtain data from 1986q2.

Table B.3: Lag reduction for the exogenous variables in the unrestricted VAR [a,]

Lags	log likelihood	SC	HQ	AIC
5	869.13433	-14.194	-15.824	-16.926
4	866.47195	-14.433	-15.964	-16.999
3	860.07987	-14.590	-16.022	-16.991
2	857.56754	-14.832	-16.166	-17.067
1	854.16023	-15.055	-16.290	**-17.124**
0	845.28489	**-15.157**	**-16.293**	-17.061

Tests of lag reduction		
5 to 4	F(6,112) =	0.55420 [0.7658]
5 to 3	F(12,148) =	0.96638 [0.4836]
5 to 2	F(18,158) =	0.83006 [0.6629]
5 to 1	F(24,163) =	0.81618 [0.7127]
5 to 0	F(30,165) =	1.0756 [0.3722]
4 to 3	F(6,116) =	1.4069 [0.2178]
4 to 2	F(12,153) =	0.98362 [0.4670]
4 to 1	F(18,164) =	0.91767 [0.5582]
4 to 0	F(24,168) =	1.2251 [0.2269]
3 to 2	F(6,120) =	0.55985 [0.7615]
3 to 1	F(12,159) =	0.66799 [0.7801]
3 to 0	F(18,170) =	1.1519 [0.3071]
2 to 1	F(6,124) =	0.78849 [0.5806]
2 to 0	F(12,164) =	1.4710 [0.1398]
1 to 0	F(6,128) =	2.1855[0.0485]*

Estimation period: 1986q2-2008q4

[a] Endogenous variables: Real housing prices, real household debt and real disposable income. Restricted variables: Real interest rate after tax, housing turnover, housing stock and a linear trend. Unrestricted variables: Constant and seasonal dummies.