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Abstract:
The financial crisis has brought the interaction between housing prices and household borrowing into the limelight of economic policy debate. This paper examines the nexus of housing prices and credit in Norway within a structural vector equilibrium correcting model (SVECM) over the period 1986q2-2008q4. The results establish a two-way interaction in the long-run, so that higher housing prices lead to a credit expansion, which in turn puts an upward pressure on housing prices. Interest rates influence housing prices indirectly through the credit channel. Furthermore, households' expectations about future development in teir own income as well as in the Norwegian economy have a significant impact on housing price growth. Dynamic simulations show how shocks are propagated and amplified. When we augment the model to include the supply side, these effects are dampened. The paper is an extended version of Anundsen and Jansen (2013b) and it encompasses a previous Discussion Paper 651 (Anundsen and Jansen, 2011).

Keywords: Housing prices, household borrowing, financial accelerator, dynamic simulations.

JEL classification: C32, C52, E27, E44, G21, G28, R21, R31.

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Sammendrag

Samspillet mellom boligpriser og husholdningenes gjeld blir belyst ved hjelp av simultan modellering både på kort og på lang sikt. Den langsiktige sammenhengen mellom de to variablene analyseres innenfor rammen av en kointegrert vektorautoregressiv modell. Realboligprisen, husholdningenes realdisponible inntekt og husholdningenes realgjeld blir forklart i denne modellen, mens realrenten etter skatt, antallet boligtransaksjoner og boligkapitalen inngår som betingingsvariable. Forfatterne identifiserer to likevektsammenhenger, som styrer henholdsvis boligpriser og husholdningenes gjeld på lang sikt. De finner at realboligprisene avhenger av husholdningenes realgjeld, realdisponibel inntekt og boligkapital i faste priser, mens realgjelden på lang sikt er bestemt av realverdien av boliger, realrenten etter skatt og antall boligtransaksjon. Dette innebærer at det er en gjensidig avhengighet på lang sikt mellom boligpriser og gjeld.

Disse langsiktssammenhengene bygges inn i et system med to likevektsjusteringsrelasjoner som tallfestes simultant på kvartalsdata for perioden 1986(2)-2008(4). Modellen viser at gjelden påvirker boligprisene direkte også på kort sikt, mens boligprisene bare påvirker gjelden indirekte via likevektsjusteringsleddet. I tillegg finner forfatterne at en forventningsvariabel, som måler husholdningenes forventninger om utviklingen i egen økonomi så vel som i makroøkonomien framover, har en klar effekt på boligprisene. Prognoseegenskapene til modellen forbedres ved å også ta med endringer i inflasjonsraten. Ved å utsette modellen for sjokk, viser de at det er klare selvforsterkende effekter mellom boligprisene og husholdningenes gjeld.

Denne publikasjonen er en utvidet versjon av Anundsen og Jansen (2013b) og den erstatter et tidligere utgitt Discussion Paper 651 (Anundsen og Jansen, 2011),

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. 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 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 the long-run equilibrium will occur more frequently in the housing market than in the financial markets.

In the housing market, 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 motivate the banks to expand their lending.

That said, most housing purchases are financed by credit, and changes in household borrowing are expected to affect housing prices. The potential self-reinforcing mechanism that works between 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 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 analyze the interaction between housing prices and credit in Norway.

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

The paper contributes to the literature in several ways. First, we use a system based cointegration analysis, while most existing studies rely on single-equation methods. We expect to find (at least) two cointegrating vectors and the system analysis is important for both identification and for estimation efficiency. The disposable income for the household sector is included as a third endogenous variable in the VAR and is found to be weakly exogenous with respect to the long-run coefficients in the model. This motivates why we focus on housing prices and credit in modeling the short-run adjustments.

Second, the dynamic interaction between housing prices and credit is also analyzed using system methods. Full information maximum likelihood is used in the design of the short-run specifications, which is carried out general-to-specific. Previous studies have resorted to an equation-by-equation approach at this stage.

Third, the paper includes a measure of households expectations about the future development in their own as well as the Norwegian economy in the dynamic specification. As a housing purchase is a long term investment, this seems to be a highly relevant variable to include in a housing price equation. Indeed, it is shown that this variable has a positive and significant impact on housing prices.

While many previous studies have had difficulties measuring supply side effects, our results indicate a large and negative long run impact on housing prices of an increase in the housing stock. This suggests that supply side constraints are important for long-run movements in prices and that a liberalization of zoning regulations and other regulations limiting the supply of housing might be an effective tool to prevent a rapid increase in housing prices.

Finally, dynamic simulations demonstrate how shocks are propagated and amplified across the two markets over time. When we take the analysis one step ahead and include a separate model for the supply side, the effects of a positive shock to housing prices or to credit are dampened over time as residential investments gradually shift the supply of housing.

The paper gives a survey of the recent literature in Section 2. A description of the Norwegian housing and credit markets is outlined in Section 3. Section 4 provides a brief theory discussion, while we investigate the fundamental determinants of housing prices and household debt in Section 5 by means of a system based cointegration analysis. Section 6 describes the dynamic interaction between the two variables. The model yields meaningful short and long term effects when estimated on the sample 1986q2-2008q4. In Section 7, we compare our basic model for housing prices and household debt with an enlarged version which also includes the supply of housing. In both cases, dynamic simulations demonstrate that there are self-reinforcing feedback effects between the two variables of interest. Before concluding, Section 8 explores the robustness and stability of the model by adding four more years of data that have become available after the model was first documented.

2 A survey of empirical contributions

The empirical literature on housing prices is extensive; see *e.g.* Hendry (1984), Muellbauer and Murphy (1997), Pain and Westaway (1997), Meen (2001, 2002) and Malpezzi (1999) to mention a few important contributions. Girouard et al. (2006) provide a nice overview of the empirical literature. The majority of the papers have investigated the determinants of housing prices within a single-equation set-up. That framework does not shed light on the possible interaction between housing prices and household borrowing. Only recently – in the past decade – a literature on the nexus of housing prices and credit has emerged. The results up to now disagree about the direction of causality. The discrepancies can, however, be ascribed to a number of sources: There are institutional differences between countries, and the methodological approaches as well as sample sizes and data sets vary across the studies. A summary of the empirical findings on the interaction between housing prices (ph) and credit (d), which we refer to below, is given in Table 1 and Table 2.

Table 1: Literature Evidence on the Long-Run Interaction Between Housing Prices and Credit[a]

Author(s)	$ph \rightarrow d$	$ph \leftarrow d$	$ph \leftrightarrow d$
Hofmann (2003, 2004)	*		
Brissimis and Vlassopoulos (2009)	*		
Gerlach and Peng (2005)	*		
Oikarinen (2009a,b)		*	
Fitzpatrick and McQuinn (2007)			*
Berlinghieri (2010)			*
Gimeno and Martinez-Carrascal (2010)			*

[a] The table summarizes the literature evidence on the long-run interaction between housing prices and credit. Housing prices are denoted by ph, while credit is denoted by d.

Table 2: Literature Evidence on the Short-Run Interaction Between Housing Prices and Credit[a]

Author(s)	$ph \rightarrow d$	$ph \leftarrow d$	$ph \leftrightarrow d$
Hofmann (2003)			*
Brissimis and Vlassopoulos (2009)			*
Gerlach and Peng (2005)	*		
Oikarinen (2009a,b)[b]		*	
Fitzpatrick and McQuinn (2007)		*	
Berlinghieri (2010)			*

[a] The table summarizes the literature evidence on the short-run interaction between housing prices and credit. Housing prices are denoted by ph, while credit is denoted by d.

[b] The results apply to the period after the Finnish credit markets were deregulated.

In an early study, using both panel data and time series techniques for 20 countries, Hofmann (2003) finds a cointegrating relationship between property prices, bank lending and GDP. The equation is interpreted as a credit equation and property prices are found to affect private sector borrowing in the long-run, while the opposite direction of causation is not supported. The data are quarterly and cover the period 1985-2001. The author also reports results for the short-run dynamics, where he finds causality to go in both directions. The long-run results are further corroborated in Hofmann (2004),[2] where he first studies 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 for each country. For a majority of the countries, the Johansen analysis (Johansen (1988)) shows no cointegration with this information set. When he extends 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 a credit equation for those countries where a high share of loans are secured by real estate.

This finding is supported by Brissimis and Vlassopoulos (2009) in a single country study for Greece. With quarterly data specific to the housing market for the period 1993-2005, they find only one cointegrating relationship based on system based cointegration techniques. This is interpreted as a mortgage loan equation, where loans are determined by housing prices, interest rates and an income measure. The loadings reveal that only the credit variable equilibrium corrects, *i.e.* housing prices are found to be weakly exogenous with respect to the long-run parameters. 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. They use a vector equilibrium correction framework and find that the direction of causation is from housing prices to private sector debt both in the long-run and in the short-run.

Contrary to this, Oikarinen (2009b) finds the direction of causation to go from household borrowing to housing prices in the long-run. He uses quarterly data for Finland from 1975 to 2006 to explore the mutual dependence between housing prices and borrowing. A cointegration analysis in the spirit of Johansen (1988) supports the existence of only one cointegrating vector, which is interpreted as a housing price equation. Tests for Granger non-causality show that there is no dynamic effect going in either direction before 1988, *i.e.* before the Finnish credit market was considered fully deregulated. There is however an effect on housing prices from the credit market running via the equilibrium correction term. After the deregulation, however, lending is shown to Granger cause housing prices also through the short-run dynamics, while the opposite is not found to be the case. Furthermore, both variables are affected by the equilibrium correction term in the short-run after the deregulation has taken place. These results are corroborated by an impulse response analysis, where Oikarinen establishes an interaction between housing prices and credit only after the deregulation process was considered completed (after 1987). Using the same methodological framework, Oikarinen (2009a) reports similar results with regional housing price data for the Helsinki Metropolitan area. Again, household debt enters the long-run relationship for housing prices and Granger non-causality tests give the same results as in Oikarinen (2009b).

There are also a few recent studies documenting a mutual dependency in the long-run, *i.e.* two cointegrating vectors are found. Fitzpatrick and McQuinn (2007) look at the interaction between housing prices and mortgage credit in Ireland between 1981

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

and 1999. They show that the two variables are mutually dependent in the long-run, as well as in the short-run. In the dynamic specification, a contemporaneous effect is only established from credit to housing prices, while housing prices are found to have lagged effects on credit. Like Hofmann (2003), Fitzpatrick and McQuinn (2007) analyze the long-run dependence within a single-equation framework adopting the original approach to cointegration of Engle and Granger (1987).[3]

When exploring the dynamic interaction between housing prices and credit, the two equations are estimated separately by OLS and a general-to-specific procedure is followed to find a parsimonious system. Acknowledging the potential endogeneity problems, Fitzpatrick and McQuinn estimate the two equations jointly by non linear three stage least squares after having sequentially reduced the dimensionality of the two equations.[4]

The results of Fitzpatrick and McQuinn (2007) are supported by Berlinghieri (2010) for quarterly US data covering the period 1977 to 2005 who also finds a bi-directional interdependence in the long-run. A two step Engle-Granger approach is adopted and the short-run dynamics are estimated by single-equation OLS. The interaction is found to run in both directions also in the short term.

Making use of quarterly data for the period 1984-2009, Gimeno and Martinez-Carrascal (2010) study the interaction between housing prices and household borrowing in Spain. A system based cointegration analysis shows that the two variables are interdependent in the long-run, *i.e.* 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 equilibrium correct to disequilibrium constellations in the housing market. They do not report results for the short-run dynamics.

An alternative approach to modeling housing prices is adopted by Carrington and Madsen (2011), who consider a Tobin's Q model for US housing price determination over the sample 1967q2-2010q2. They use an ARDL bounds testing approach to test whether housing prices, the cost of agricultural land and construction costs are cointegrated. They do not find evidence for cointegration and consider a model in first differences instead. Interestingly, they find an important role of banks' willingness to lend for short-run fluctuations in housing prices. These results are confirmed by a panel analysis for eight OECD countries over the period 2003q1–2010q3.

The diverging results, as summarized in Table 1 and Table 2 call for further research. Our paper adopts the same econometric approach as Gimeno and Martinez-Carrascal (2010), but we go further. Not only do we to study the long-run interaction, but also the dynamic interaction between the two markets, which is important for both policy evaluation and forecasts.

The studies that address the short-run interaction by modeling the dynamics of the two variables all use a single-equation approach, *i.e.* the equations are estimated separately by OLS regressions. In some cases, the system is estimated jointly by 3SLS after

[3]Hofmann (2003) also considers a Johansen analysis, but it is the results from the single-equation procedure that are retained for the dynamic specifications.

[4]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 hence it is not directly integrated in their analysis.

the dimensionality of the equations in the system have been reduced separately. This may be inappropriate – as pointed out by Hammersland and Jacobsen (2008) – 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. From this perspective, it seems highly relevant to deal with the potential simultaneity from the onset. Hence, one should design the structural short-run model using system methods that takes on the simultaneity problem from the outset.

3 The Norwegian housing and credit markets

The banking crisis in Norway that took place between 1988-1993 is a clear example of a collapse of property prices being followed by imbalances in the real economy. The recent financial crisis was different in that it was an external shock to the domestic economy, which had a significant, but short-lived, negative effect on Norwegian housing prices.

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.

For our purpose, it is important to note that also the housing market was heavily regulated in Norway after World War II. Building materials were rationed and there were strict regulations on housing, both with regard to quantity and prices. These regulations ended in July 1982, with the abolition of price regulation on cooperative housing. The credit market regulations were lifted shortly after this. 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 Norwegian banking crisis, which ended in 1993, real housing prices have grown almost consecutively until the financial meltdown of the previous decade (see Figure 1a). Growing housing prices have been accompanied by a substantial expansion in real household debt (see Figure 1b).

The historical episodes referred to above strongly 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 1c plots the four quarter growth in real housing prices against percentage deviations of GDP mainland Norway from trend.[5] A close link between economic activity and housing prices is apparent over the entire period, with a less pronounced correlation pattern the last few years. Goodhart and Hofmann (2007) argue that there will be 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 1c shows such a tendency for the case of Norway in the period after the deregulation of the Norwegian credit markets had been completed. 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

[5]GDP mainland Norway measures total production in Norway excluding two sectors: extraction of oil and gas, and ocean transport.

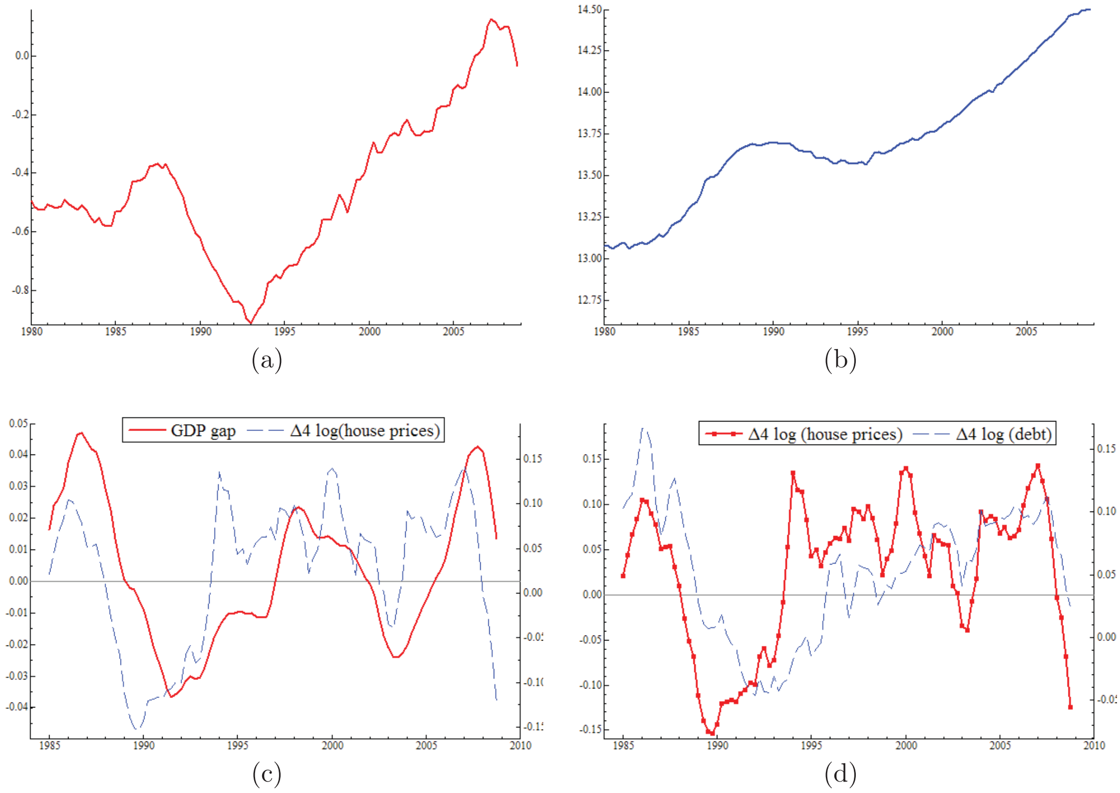


Figure 1: Panel a) Log of real housing prices, 1980-2008. Panel b) Log of real household debt, 1980-2008. Panel c) GDP gap (left scale) and four quarter growth in real housing prices (right scale), 1985-2008. Panel d) Four quarter growth in real housing prices (left scale) and in real household debt (right scale), 1985-2008.

costs, that is the Tobin q (Tobin, 1969) for residential investments. Another channel in which housing prices could have an effect on the business cycle is by amplifying shocks in the credit market. It is evident from Figure 1d, where we have plotted the four quarter growth in real housing prices against four quarter growth in real household borrowing, that the two series move quite closely together.

Previous studies of the credit and housing markets in Norway do not take the potential simultaneity between the two into account. For example, the determination of household debt is the topic of Jacobsen and Naug (2004), whilst Jacobsen and Naug (2005) describe a separate model for housing prices. In Jacobsen and Naug (2004), housing prices are one of the fundamental factors explaining household debt, whereas household borrowing is not part of the cointegrated vector explaining housing prices in Jacobsen and Naug (2005).[6] That said, it is documented that the interest rate is an important determinant of housing prices. Also, Jacobsen and Naug (2004) find that the interest rate is one of the fundamental factors explaining household borrowing. The effect of interest rates on credit thus suggests that the interest rate variable in the housing price equations captures a credit effect, *i.e.* the coefficient of the interest rate in Jacobsen and Naug (2005) picks

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

up a gross effect.[7]

4 Economic theory

The commonly used framework for modeling housing prices is the life-cycle model, see *e.g.* Meen (2001, 2002), Muellbauer and Murphy (1997, 2008) and the references therein. We augment this model with a term capturing the presence of credit constraints, and the marginal rate of substitution (MRS) between housing and a composite consumption good is then given by (see *e.g.* Meen (1990) or Meen and Andrew (1998)):

$$MRS = PH_t \left[(1-\tau_t)i_t - \pi_t + \delta_t - \frac{\dot{PH}_t^e}{PH_t} + \lambda_t/\mu_c \right], \tag{1}$$

where PH_t is real housing prices, τ_t is the marginal tax rate on equity income, i_t is the nominal interest rate (paid by households for loans), π_t is the annual inflation rate, δ_t is the depreciation rate or the rate of maintenance costs including property taxation, and $\frac{\dot{PH}_t^e}{PH_t}$ is the expected real rate of appreciation for housing prices. λ_t is the shadow price of the credit constraint which is divided by the marginal utility of consumption μ_c. This is commonly known as the real housing user cost of capital, in this case augmented with a credit constraint. Market efficiency requires that the following no-arbitrage relationship holds, where Q_t represents the real imputed rental price for housing services

$$PH_t = \frac{Q_t}{(1-\tau_t)i_t - \pi_t + \delta_t - \frac{\dot{PH}_t^e}{PH_t} + \lambda_t/\mu_c} \tag{2}$$

Meen (2002) follows Poterba (1984) and interprets (2) as an inverted housing stock demand function. In the following, we will assume that the depreciation rate is constant.[8] If we assume that Q_t, which is unobservable, is a function of real disposable income for the household sector (excluding dividends), YH_t, and the stock of dwellings, H_t, we can write the inverted demand function as

$$PH_t = f^* \left(H_t, YH_t, R_{t,} \frac{\dot{PH}_t^e}{PH_t}, \lambda_t/\mu_c \right), \tag{3}$$

where $R_{t,}$ is the real after tax interest rate $(1-\tau)i_t - \pi$.

With a constant depreciation rate, the real user cost can be split in two different components: The real direct user cost (as measured by R_t) and expected real housing price appreciation. In the econometric analysis, we use the real direct user cost as our operational measure of the user cost and let price expectations be modeled by allowing

[7]Akram et al. (2006), Akram et al. (2007) and Andersen (2011) augment the core part of a macroeconometric model for the Norwegian economy (see e.g. Bårdsen et al. (2003) and Bårdsen et al. (2005)) with different versions of the housing price and credit equations of Jacobsen and Naug (2004, 2005). These studies address issues related to financial stability when there are interaction effects between housing prices and credit.

[8]Assuming a constant depreciation rate is consistent with the Norwegian National accounts, where a constant depreciation rate is used for housing.

lagged real price appreciation to enter our dynamic model.[9] This is similar to Abraham and Hendershott (1996), Gallin (2008) and Anundsen (2012) on US data, and it is consistent with the lagged housing price appreciation not having permanent effects, but rather that it picks up a momentum or the "bubble builder" effect using the terminology of Abraham and Hendershott (1996).[10]

Furthermore, we shall substitute household loans as a proxy for the theoretically correct – but unobservable – λ_t/μ_c term in (3).[11] Our empirical study can thus be seen as a test of the informational value of household loans when direct information on credit constraints is missing. As household debt is non-stationary, we implicitly assume that the same holds for the shadow price of the credit constraint.

Hence, we formulate the determination of real housing prices at the aggregate level in a static long-run equilibrium as

$$PH_t = f(H_t, YH_t, R_t, D_t), \tag{4}$$

where $\frac{\partial f}{\partial H} < 0, \frac{\partial f}{\partial YH} > 0, \frac{\partial f}{\partial R} \gtrless 0, \frac{\partial f}{\partial D} > 0$ and D_t is real household debt.

Equation (4) expresses market clearing prices for any given level of the housing stock. The equation 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. The sign of the derivative with respect to the interest rate is ambiguous. The main effects of a change in the interest rate work through disposable income and household loans, which both are controlled for in (4). What remains are the substitution effects which may be of either sign from a theoretical point of view.[12]

We supplement our model for housing prices with a relationship that determines real household debt in a long-run equilibrium

$$D_t = g(H_t, YH_t, R_t, PH_t, TH_t), \tag{5}$$

where $\frac{\partial g}{\partial H} > 0, \frac{\partial g}{\partial YH} > 0, \frac{\partial g}{\partial R} < 0, \frac{\partial g}{\partial PH} > 0, \frac{\partial g}{\partial TH} > 0$ and TH_t denotes the housing turnover. Equation (5) 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. Since all the variables

[9] It should be mentioned that we have experimented with a moving average process for the expectation component of the user cost. We find that this term is insignificant in our long-run relationships, suggesting that it is reasonable to assume that lagged price appreciation effects are picked up through the dynamics of the model. We then avoid making *a priori* assumptions about the expectation formation.

[10] Abraham and Hendershott (1996) distinguish between a bubble builder effect represented by lagged real housing price appreciation in the dynamic part of the model and a bubble burster effect through an equilibrium correction term.

[11] An alternative approach has been considered in Duca et al. (2011a,b) on US data. Including a measure of the LTV ratio for first time home buyers, they find that exogenous shifts in credit conditions have been important for US housing price dynamics in the 2000s.

[12] It is not only from a theoretical point of view that the sign of the direct effect is ambiguous. Empirically it is often found to be statistically insignificant. In the case of Norway the dominant interest rate effects on housing prices are indirect. Almost all mortgage debt in Norway are loans with flexible interest rates. Hence, 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. The inclusion of the credit aggregate captures the effect on housing prices from a change in the cost of financing.

included in (4) and (5) are usually found to be non-stationary and integrated of first order, and since the theory postulates long-run equilibrium relationships, the discussion in this section suggests that housing prices and credit should be cointegrated with the variables – or a subset thereof – included in (4) and (5), i.e. we would expect to find two cointegrating relationships.

In the following we shall think of equations (4) and (5) as a subsystem, conditioning on H_t, YH_t, R_t, and TH_t. The last three variables can be assumed to be determined by factors other than housing prices and credit. The housing stock, H_t, on the other hand represents the supply side of the housing market. It appears in equation (3) since it affects negatively the market clearing rent and hence the price of housing. We will assume it is related to the profitability of new construction and thus that it is influenced positively by real housing prices and negatively by construction costs. Hence, there are feedback effects from housing prices via H_t to housing prices and credit. In order to capture these feedback effects we estimate a submodel for housing supply separately in Appendix A. In Section 7, when we compare the dynamic responses from our baseline model with those from an extended version of the model, which includes the housing supply, we find that the effects of a shock to housing prices or household debt are dampened.

5 Cointegration analysis

5.1 Methodological approach

A semi-logarithmic transformation of the variables appearing in equations (4) and (5) – which can be seen as a linearization of the theoretical formulations – forms the basis for the information set underlying our empirical analysis. All data are seasonally unadjusted and in what follows, small letters indicate that the variables are measured on a logarithmic scale.[13] All monetary variables are measured in real terms, having been deflated by the consumption deflator. Our sample covers the period 1986q2-2008q4. We have data for the number of housing transactions only from 1985q1, and the housing price data are also less reliable in the period prior to this. Since we consider a post-deregulation sample, it follows that we do not account for shifts in the constraints that are due to the deregulation of the Norwegian housing and credit markets. That said, the deregulation of the housing and credit markets in the early 1980's is likely to have altered the functioning of both, so that a different econometric model would probably be more suitable if we were to consider the period prior to the deregulation. In particular, it is less likely that a self-reinforcing relationship between housing prices and credit existed during the regulation period, since these regulations clearly distorted the ordinary market mechanisms.[14]

The orders of integration of the data series have been examined by a suite of different tests; the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller (1979)), the Phillips-Perron (PP) test (Phillips (1987) and Phillips and Perron (1988)), as well as

[13] For a detailed data description, see Appendix B. The log transformation is applied to all variables in (4) and (5), except the real after tax interest rate.

[14] This is consistent with the empirical findings of Oikarinen (2009b), who finds that a two-way interaction between housing prices and credit in Finland can only be established after liberalization of the credit markets in the late 1980s.

the Kwiatowski, Phillips, Schmidt, and Shin (KPSS) test (Kwiatkowski et al. (1992)).[15] Based on these tests, we treat all variables as integrated of order one at most in the econometric analysis. There is also supporting evidence for this approach in that we find - as we report below - that the residuals in the final empirical model turn out to be stationary. Details on the tests for unit roots are given in Table C.1 of Appendix C.

Due to the non-stationarity of the variables in our data set, we start by investigating the the long-run determinants of housing prices and household borrowing in a cointegrated VARX system where also household income is treated as an endogenous variable, while we condition on the real after tax interest rate, the housing turnover and the housing stock. Finding evidence of cointegration ensures that we can formulate the VARX as a vector equilibrium correction model (VECM). The VECM approach provides an opportunity to study long-run determinants and short-run dynamics in a unified framework, which opens for the possibility that the causality between housing prices and credit is bi-directional both in the short-run and in the long-run. The model is therefore suitable for addressing the key issue: Is there empirical evidence for the existence of a financial accelerator in the Norwegian housing market?

In general, the I(1) cointegrated VAR (CVAR) model can be written as a re-parameterization of a $VAR(p)$ model, see for example Johansen (1988), Johansen (1995) and Juselius (2006):

$$\Delta \boldsymbol{Y}_t = \boldsymbol{\Pi} \boldsymbol{Y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{Y}_{t-1} + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t, t = 1, ..., T \tag{6}$$

$\boldsymbol{Y}_t$ is a $n \times 1$ matrix comprising the endogenous variables in the system, while $\boldsymbol{D}_t$ contains deterministic terms such as a constant, linear trends or other regressors considered to be fixed. We let $\boldsymbol{\Pi}$, $\boldsymbol{\Gamma}_i$ and $\boldsymbol{\Phi}$ denote the coefficient matrices. With reference to a $VAR(p)$ model, the $\boldsymbol{\Pi}$ and $\boldsymbol{\Gamma}_i$ matrices are defined as $\boldsymbol{\Pi} = \sum_{i=1}^{p} \boldsymbol{\Pi}_i - \boldsymbol{I}$ and $\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{p} \boldsymbol{\Pi}_j$, where $\boldsymbol{\Pi}_i$ is the VAR coefficient matrix attached to lag number i. The innovation terms, $\boldsymbol{\varepsilon_t}$, are assumed to be independently Gaussian distributed, N(0,$\boldsymbol{\Sigma}$), and the initial values $Y_{-p}, ..., Y_0$ are considered fixed.

In our case, we consider a $VARX(p,q)$, i.e. some of the variables in the system are treated as weakly exogenous. In addition, we follow the suggestion of Harbo et al. (1998) for partial systems and restrict a deterministic trend to enter the cointegration space. Thus, the $VECM(p,q)$ representation of the $VARX(p,q)$ that forms the basis for our econometric analysis reads:

$$\Delta \boldsymbol{X}_t = \tilde{\boldsymbol{\Pi}} \tilde{Y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{X}_{t-i} + \sum_{i=0}^{q-1} \boldsymbol{\Psi}_i \Delta \boldsymbol{Z}_{t-i} + \tilde{\boldsymbol{\Phi}} \tilde{\boldsymbol{D}}_t + \boldsymbol{\varepsilon}_t. \tag{7}$$

where $\boldsymbol{X}_t$ is a 3×1 matrix comprising the endogenous variables ph, d and y, while $\mathbf{Y}_t = (\mathbf{X}_t', \mathbf{Z}_t')'$ is a $(3+3) \times 1$ matrix where $\mathbf{Z}_t$ is a 3×1 matrix composed of the weakly

[15] As a guidance for choosing the optimal lag truncation for the ADF test, we have relied on Akaike's information criterion (AIC) starting with an initial lag length of eight in the first differences in all test regressions and then chosen the specification with the lowest AIC value.

exogenous variables R, t and h and $\tilde{Y}_t = (\boldsymbol{Y}_t', t)'$ with t denoting a deterministic trend. The vector $\tilde{\boldsymbol{D}}_t$ comprise a constant and centered seasonal dummies.

The trace test for the order of cointegration (Johansen, 1988) can be used to determine the rank of the matrix $\tilde{\boldsymbol{\Pi}}$, which corresponds to the number of independent linear combinations between the variables that are stationary. We follow Johansen (1988) and define $\tilde{\boldsymbol{\Pi}} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\beta}$ is a $(n+k+1) \times r$ matrix and $\boldsymbol{\alpha}$ is a $n \times r$ matrix corresponding to the long run coefficients and loading factors respectively. The rank of the $\tilde{\boldsymbol{\Pi}}$ matrix is denoted by r, while n refers to number of endogenous variables and $k+1$ is the number of exogenous variables (including the deterministic trend, which is restricted to lie in the cointegration space). Thus, in our case – with $n = k = 3$ – $\boldsymbol{\beta}$ is a $7 \times r$ matrix and $\boldsymbol{\alpha}$ is a $3 \times r$ matrix.

5.2 Cointegration results

As mentioned, our starting point for the cointegration analysis is a VARX in real housing prices, real household debt and real disposable income, while we condition on the real after tax interest rate, the housing turnover and the housing stock.[16] We start with a lag length of 5 in both the endogenous and the weakly exogenous variables ($p = q = 5$), which ensures that we have a well specified model without evidence of autocorrelation, heteroskedasticity nor non-normality. Then, the optimal lag truncation is decided based on AIC. According to AIC, the VAR-model should include 5 lags in the endogenous variables, while we find that only one lag is needed for the weakly exogenous variables.[17]

Having decided on the lag length, we use the trace test to decide on the number of cointegrating relationships. Table 3 displays the results. We find that there are two cointegrating vectors.[18] The model is well specified – residual diagnostics show that the residuals are neither heteroskedastic nor autocorrelated, and normality is not rejected.

Exact identification can be achieved by imposing two restrictions in each vector. We start by normalizing on real housing prices in the first vector and real household debt in the other. In addition, it is assumed that the housing turnover has no direct effect on real housing prices.[19] This is in accordance with the theoretical housing price equation (4), while earlier studies have found that the turnover affects household borrowing in Norway (see Jacobsen and Naug (2004)), which suggests that it should be part of the relationship determining household debt. The final restriction we use for exact identification is that it is the value of the housing capital – and not simply housing prices – which determines the

[16]Indeed, including the turnover as an endogenous variable in the VAR, we find that it is weakly exogenous (the p-value from the test is 0.6847). This supports our conditioning and saves valuable degrees of freedom. Alternatively, weak exogeneity can be tested along the lines of Johansen (1992), Harbo et al. (1998), Pesaran et al. (2004) and Dees et al. (2007), i.e. by including the two cointegrating vectors we document below in the marginal model for the turnover and then test their joint significance. An F-test of the two zero restrictions has a p-value of 0.1891, which gives further justification to this assumption.

[17]Details are available in Table C.2 in Appendix C.

[18]Critical values correcting for the inclusion of exogenous variables (see Doornik (2003)) have been used.

[19]Gimeno and Martinez-Carrascal (2010) and Fitzpatrick and McQuinn (2007) exclude the real interest rate from the long-run equation for housing prices by assumption. Pursuing this alternative identification strategy, i.e. excluding the real interest rate instead of the turnover from the housing price equation from the outset, we get identical results to those reported below.

Table 3: Trace test for cointegration [a]

$Eigenvalue : \lambda_i$	H_0	H_A	λ_{trace}	5%-critical value[b]
0.39	$r = 0$	$r \geq 1$	86.59	64.48
0.22	$r \leq 1$	$r \geq 2$	41.74	40.95
0.19	$r \leq 2$	$r \geq 3$	18.82	20.89
Diagnostics[c]	Test statistic	Value[p-value]		
Vector AR 1-5 test:	F(45,146)	1.06 [0.39]		
Vector Normality test:	$\chi^2(6)$	7.78 [0.26]		
Vector Hetero test:	F(270,247)	1.03 [0.42]		
Estimation period:	1986q2-2008q4			

[a] Endogenous variables: Real housing prices (ph), real household debt (d) and real disposable income (yh). Restricted variables: Real interest rate after tax (R), housing turnover (th), housing stock (h) and a trend (t). 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 (2009a).

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.

Based on the identified cointegrated vectors, we can move on to test overidentifying restrictions. The results for these restrictions are documented in Table 4 below.[20] For every new restriction that is imposed, we report both the log-likelihood value, the incremental test as well as the total test at the bottom line of each panel. In Panel 1, the trend variable is dropped from both equations, which correspond to two testable overidentifying restrictions. Next, in Panel 2, we omit the real after tax interest rate from the vector associated with real housing prices. As mentioned above, this does not imply that a change in the interest rate will not affect housing prices, but it means that interest rate effects are captured by changes in disposable income and through the credit channel. In Panel 3, there is no effect of disequilibrium in the housing market on household debt, whereas Panel 4 shows the case with no direct effect of real disposable income on household debt. Finally, Panel 5 shows the result when we impose that the loadings of both cointegrating vectors with respect to income are zero, *i.e.* the test shows weak exogeneity of income with respect to the long-run coefficients, see Johansen (1992). According to the incremental tests reported in Table 4, all 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 4, describe the two final long-run relationships for housing prices and household debt.[21] 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 in the final model (Panel 5) and they are significant at conventional significance levels.[22]

[20] The absolute value of standard errors are reported in parentheses below the estimated coefficients.

[21] In Table C.3 in Appendix C, we report the loading factors corresponding to each of the panels.

[22] The interest rate is the only exception. However, using a one sided test, which appears to be 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 4: Testing steady-state hypotheses.

The just identified house price and debt equations are defined by

$$ph = \beta_{d,1}d + \beta_{yh,1}yh + \beta_{h,1}h + \beta_{R,1}R + \beta_{t,1}t$$
$$d = \beta_{ph,2}ph + \beta_{yh,2}yh + \beta_{R,2}R + \beta_{th,2}th + \beta_{h,2}h + \beta_{t,2}t$$

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

$$ph = \underset{(0.07)}{0.76}d + \underset{(0.21)}{1.39}yh - \underset{(0.37)}{2.00}h + \underset{(0.85)}{0.13}R$$
$$d = 1.53ph - \underset{(0.17)}{1.45}yh - \underset{(1.40)}{0.71}R + \underset{(0.05)}{0.09}th + \underset{(0.07)}{1.53}h$$
$$LogL = 842.845\ , \quad \chi^2(2) = 3.81[0.15]$$

Panel 2: No effect of real after tax interest rate on house prices ($\beta_{R,1} = 0$)

$$ph = \underset{(0.08)}{0.77}d + \underset{(0.22)}{1.43}yh - \underset{(0.40)}{2.07}h$$
$$d = 1.54ph - \underset{(0.18)}{1.48}yh - \underset{(0.40)}{0.54}R + \underset{(0.05)}{0.10}th + \underset{(0.07)}{1.54}h$$
$$LogL = 842.834\ , \ \chi^2(1) = 0.02[0.88], \ \chi^2(3) = 3.84[0.28]$$

Panel 3: No effect of disequilibrium housing prices on household debt

$$ph = \underset{(0.19)}{0.84}d + \underset{(0.65)}{1.67}yh - \underset{(1.18)}{2.58}h$$
$$d = 1.08ph - \underset{(0.85)}{1.18}yh - \underset{(2.35)}{3.98}R + \underset{(0.28)}{0.56}th + \underset{(0.30)}{1.08}h$$
$$LogL = 842.276\ , \ \chi^2(1) = 1.12[0.29], \ \chi^2(4) = 4.95[0.29]$$

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

$$ph = \underset{(0.19)}{0.86}d + \underset{(0.64)}{1.42}yh - \underset{(1.16)}{2.33}h$$
$$d = 0.78ph - \underset{(1.87)}{2.83}R + \underset{(0.15)}{0.24}th + \underset{(0.15)}{0.78}h$$
$$\text{LogL} = 841.323\ , \ \chi^2(1) = 1.12[0.29], \ \chi^2(5) = 6.86[0.23]$$

Panel 5: Imposing weak exogeneity of income with respect to the long-run coefficients :

$$ph = \underset{(0.19)}{0.98}d + \underset{(0.63)}{1.69}yh - \underset{(1.15)}{3.03}h$$
$$d = 0.76ph - \underset{(1.79)}{2.74}R + \underset{(0.15)}{0.28}th + \underset{(0.16)}{0.76}h$$
$$\alpha_{1,ph} = \underset{(0.04)}{-0.24}, \alpha_{1,d} = \underset{(0.03)}{-0.10}, \alpha_{2,d} = \underset{(0.01)}{-0.04}$$
$$LogL = 840.529\ , \ \chi^2(2) = 1.59[0.451], \ \chi^2(7) = 8.44[0.30]$$

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

Note: For notation, confer footnote a in Table 3 and the variable definitions in Appendix B.

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 Martinez-Carrascal (2010) who consider nominal instead of real interest rates. It is however greater than the estimates found by Brissimis and Vlassopoulos (2009) for Greece and Fitzpatrick and McQuinn (2007) for Ireland who both consider 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 Martinez-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 *percent* in the long-run.

The adjustment coefficients (confer Panel 5) imply that both housing prices and household debt equilibrium correct when the latter departs from the value implied by its fundamentals ($\alpha_{1,d} = -0.1$ and $\alpha_{2,d} = -0.04$). Moreover, the analysis indicates that only housing prices equilibrium correct when housing prices are deviating from their steady state level ($\alpha_{1,ph} = -0.24$). This result is supported by Gimeno and Martinez-Carrascal (2010) for the case of Spain. It is interesting to note that housing prices are adjusting more rapidly to equilibrium than household debt. This is because the volume of debt is not that easily changed over night.[23]

It is worth emphasizing that our results does not suggest any separate population effects on neither housing prices nor household borrowing. This can easily be seen by reparameterizing the two cointegrating relationships in per capita terms.

$$ph = \beta_{d,1}\frac{d}{pop} + \beta_{yh,1}\frac{yh}{pop} + \beta_{h,1}\frac{h}{pop} + (\beta_{d,1} + \beta_{yh,1} + \beta_{h,1})\, pop$$
$$d = \beta_{ph,2}ph + \beta_{R,2}R + \beta_{th,2}th + \beta_{h,2}\frac{h}{pop} + (\beta_{ph,2} - 1)\, pop$$

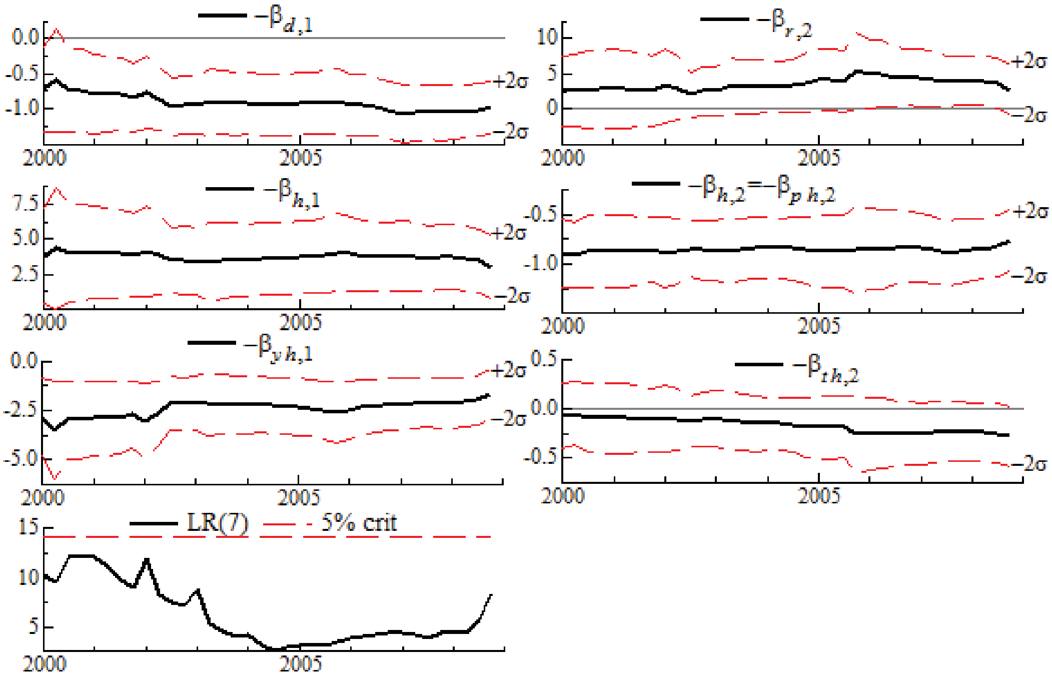
where pop is log population. Thus, for the model to imply no additional population effects, the two additional restrictions that $\beta_{d,1} + \beta_{yh,1} + \beta_{h,1} = 0$ and $\beta_{ph,2} = \beta_{h,2} = 1$ needs to hold. Imposing these two restrictions gives a p-value of 0.2449 for all nine restrictions imposed on the system, while the partial test for the two restrictions has a p-value of 0.2203. Thus, we can conclude that there is no loss of generality from not including a separate population variable in the model, which save us valuable degrees of freedom.

To investigate the recursive stability of the two long-run relationships, we have estimated the model quarter-by-quarter over the period 2000q1–2008q4. The recursively estimated coefficients are shown in Figure 2. It is clear that all the long-run coefficients in both vectors are fairly stable when estimated recursively. The lower left panel shows the recursively estimated likelihood ratio statistic[24] against the 5% critical value from the χ^2 distribution, and we see that the restrictions are accepted recursively as well.

[23] While we have only reported the adjustment coefficients from the final long-run relationships in Table 4, Table C.3 in Appendix C reports the adjustment coefficients corresponding to all the panels in Table 4.

[24] The unrestricted likelihood ($LogL_{UR}$) is derived from the model in Panel 1, while the restricted likelihood ($LogL_R$) is based on the model reported in Panel 5. The likelihood ratio statistic is then calculated as $-2\,(Lik_R - Lik_{UR})$.

Figure 2: Recursively estimated coefficients and likelihood ratio test, 2000q1–2008q4



6 Short-run dynamics

6.1 Methodological approach

To derive the simultaneous equation system, the structural vector equilibrium correction model (SVECM), that forms the basis for the analysis of the short-run dynamics, we premultiply the reduced form representation in (7) by the (non-zero) contemporaneous feedback matrix, $\mathbf{B}$:

$$\mathbf{B}\Delta\mathbf{X}_t = \mathbf{B}\tilde{\mathbf{\Pi}}\tilde{\mathbf{Y}}_{t-1} + \sum_{i=1}^{4}\mathbf{B}\mathbf{\Gamma}_i\Delta\mathbf{X}_{t-i} + \sum_{i=0}^{4}\mathbf{B}\mathbf{\Psi}_i\Delta\mathbf{Z}_{t-i} + \boldsymbol{B}\boldsymbol{\Phi}\boldsymbol{D}_t + \mathbf{B}\epsilon_t \qquad (8)$$

where we now define $\mathbf{B}\tilde{\mathbf{\Pi}} = \boldsymbol{B}\boldsymbol{\alpha}\boldsymbol{\beta}' = \boldsymbol{\alpha}^*\boldsymbol{\beta}', \boldsymbol{B}\mathbf{\Gamma}_i = \mathbf{\Gamma}_i^*, \boldsymbol{B}\boldsymbol{\Psi}_i = \mathbf{\Psi}_i^*, \boldsymbol{B}\boldsymbol{\Phi} = \mathbf{\Phi}^*, \boldsymbol{B}\boldsymbol{\epsilon}_t = \boldsymbol{\varepsilon}_t$. The new error term will also be IIN with zero mean and variance-covariance matrix given by: $\boldsymbol{\Omega} = E(\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t') = \mathbf{B}E(\boldsymbol{\epsilon}_t\boldsymbol{\epsilon}_t')\mathbf{B}' = \mathbf{B}\mathbf{\Sigma}\mathbf{B}'$.

As the income variable was found to be weakly exogenous, we can write the above system as a conditional system for housing prices and credit and a marginal model for income (see e.g Johansen (1992)). Since the focus of our paper is the interaction between housing prices and credit, we can, without loss of generality, abstract from modeling the marginal model for income. In that case, the conditional SVECM takes the following form:

$$\Delta ph_t - b_{12}\Delta d_t = \sum_{i=1}^{4} \mathbf{\Gamma}^*_{1i}\Delta\mathbf{X}^*_{t-i} + \sum_{i=0}^{4} \mathbf{\Psi}^*_{1i}\Delta\mathbf{Z}^*_{t-i} + \sum_{i=1}^{4} \widetilde{\Psi}_{1,R_i}\Delta R_{t-i} + \mathbf{\Phi}^*_1\boldsymbol{D}_t + \alpha^*_{1,ph}ECM^{ph}_{t-1} + \alpha^*_{1,d}ECM^{d}_{t-1} + \varepsilon_{ph,t} \tag{9}$$

$$-b_{21}\Delta ph_t + \Delta d_t = \sum_{i=1}^{4} \mathbf{\Gamma}^*_{2i}\Delta\mathbf{X}^*_{t-i} + \sum_{i=0}^{4} \mathbf{\Psi}^*_{2i}\Delta\mathbf{X}^*_{t-i} + \sum_{i=1}^{4} \widetilde{\Psi}_{2,R_i}\Delta R_{t-i} + \mathbf{\Phi}^*_2\boldsymbol{D}_t + \alpha^*_{2,ph}ECM^{ph}_{t-1} + \alpha^*_{2,d}ECM^{d}_{t-1} + \varepsilon_{d,t} \tag{10}$$

where we have normalized such that the contemporaneous feedback matrix, $\mathbf{B}$, has ones along the main diagonal. $\mathbf{X}^*_t$ now consists of the two remaining endogenous variables, while $\mathbf{Z}^*_t$ still represents a vector of the weakly exogenous variables in the system (including the income variable). The constant and the centered seasonal dummies are collected in $\boldsymbol{D}_t$. $\mathbf{\Gamma}^*_{ji}$, $\mathbf{\Psi}^*_{ji}$, $\widetilde{\Psi}_{j,R_i}$ and Φ^*_j (j=1, 2) are the short run coefficients, where $\mathbf{\Gamma}^*_i = (\mathbf{\Gamma}^*_{1i}, \mathbf{\Gamma}^*_{2i})$,$\mathbf{\Psi}^*_i = (\mathbf{\Psi}^*_{1i}, \mathbf{\Psi}^*_{2i})$ and $\mathbf{\Phi}^* = (\mathbf{\Phi}^*_1, \mathbf{\Phi}^*_2)$. Since the housing stock adjusts slowly, it is assumed to be fixed in the short run and is not part of the vector $\mathbf{Z}^*_t$. Note also that we have excluded the contemporaneous value of the change in real after-tax interest rate, ΔR_t, from both equations to form our general unrestricted model. However, we supplement the short run dynamics by including an expectations variable, E, which measures households expectations about future developments in their personal economy and the macroeconomy. Hence, $\mathbf{Z}^*_t$= (th, E, yh). This is the system that constitutes the general unrestricted model. This variable can also be considered as a proxy for the expected rate of appreciation in housing prices, cf. Section 4.[25]

6.2 Results for dynamic model

The simultaneous equation system represented by (9) and (10) is estimated and designed simultaneously, and once again we have to face the tough and non-trivial decision of how to exactly identify the system. To achieve exact identification, we have chosen to exclude the contemporaneous effect of the turnover in the housing price equation, while the credit equation is identified by omitting the contemporaneous value of the expectations variable. The just identified system is estimated by FIML (full information maximum likelihood). The resulting model produces well behaved residuals and serves as a starting point for the reduction process to obtain a parsimonious representation of the system.

A parsimonious model is found by stepwise elimination of insignificant variables in the system, which are excluded either one by one or in blocks. Unlike the single-equation case, no algorithm for automatic general-to-specific search exists as yet, so we have carried out the search manually.[26] In that process, we make sure that, according to the diagnostic tests, the Gaussian properties of the residuals are retained and that all imposed restrictions are supported by the data. In the preferred (final) model, we have

[25]The expectations variable is only available 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.

[26]See Doornik (2009) for a description of the automatic specification search in the case of a single-equation.

chosen to retain the income variable in the credit equation, which is relevant from *a priori* theoretical considerations, although it should have been excluded at the early stages of the reduction process had we followed a strict general-to-specific procedure. By doing so, we have achieved a more theoretically and intuitively appealing model formulation than we would have obtained otherwise, *i.e* if we had systematically eliminated the most insignificant variable at each stage. This procedure of structural model design results in the specifications displayed in Table 5.[27]

Table 5: 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		$\chi^2(46) =$	55.79 [0.1528]	
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)	0.90 [0.59]	
Vector Normality test:		$\chi^2(4)$	5.34 [0.25]	
Vector hetero test:		F(183,81)	0.88 [0.76]	
Estimation Method	FIML			
Sample	1986q2-2008q4 ($T = 91$)			

[a] Absolute t-values are reported.

[b] See Doornik and Hendry (2009a).

Table 5 reveals that credit 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 though. It is however clear that the credit aggregate will be

[27] Unlike previous studies (cf. 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. Another approach, commonly used in the literature, is instead to simplify the two equations individually before estimating them as a system. Comparing our results to the results we would have obtained following this approach, we find that the methodology followed in this paper produces results that are both more reasonable and easier to interpret from an economic point of view. Details are available in Appendix D.

influenced by housing prices through the equilibrium correction term present in the credit equation. This means that it takes about one quarter before a shock to housing prices is transmitted to the credit market. Consistent with the cointegration analysis, the short-run analysis indicates that both housing prices and household debt equilibrium correct when household debt is high relative to its stable long-run equilibrium and that only housing prices equilibrium correct when departing from their fundamentals. Our results suggest that if housing prices depart from their long-run equilibrium by one percent, housing prices will fall by -0.175 *percent*. This is greater than what is found by Jacobsen and Naug (2005),[28] but lower than the estimate reported by Fitzpatrick and McQuinn (2007).

Like Jacobsen and Naug (2004, 2005) and Fitzpatrick and McQuinn (2007), we find that the credit aggregate has a slower adjustment towards equilibrium when it is departing from its fundamentals than do 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 Martinez-Carrascal (2010), however, find the opposite to be the case for Spain.

All estimated coefficients have the expected signs. Interestingly, we find that changes in expectations have a great impact on housing prices. 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. As the equilibrium correction term for household debt is present in the housing price equation, the interest rate feeds into housing prices also here. The diagnostics indicate that the model is well specified and we find support for the imposed restrictions (p-value = 0.1528). The residuals from the two estimated equations are clearly stationary (see Table C.4 in Appendix C.

In Figure 2, we have plotted *ex ante* dynamic forecasts for the two endogenous variables. The forecasts are conditional on the explanatory variables as they accrued. The model does not fare too bad in *ex ante* forecasting, with two exceptions: the model under predicts the rapid recovery of house prices in the first half of 2009. More importantly, the credit forecasts are outside the forecast confidence bands in 2010q1 and 2011q1. However, this can for a large part be attributed to extremely cold winters, which lead to an extraordinary jump in electricity prices in each of those quarters and thus affected the consumption deflator we have used for the nominal to real transformations.

To explore this formally, Figure 3 shows *ex ante* dynamic forecasts for the two variables based on a slightly modified version of the model, where we have de-restricted the short-run price homogeneity.[29] Hence, we included the change in the price deflator, contemporaneously and at the first lag, in the short-run model. While both could be excluded from the housing price equation, both were significant with opposite signs in the credit equation. In fact, we can not reject the hypothesis that the two coefficients are equal in absolute value, i.e. suggesting that this captures a surprise inflation. As is seen, the forecasting accuracy of the model is improved. The forecasts for credit growth in 2010q1 and 2011q1 are no longer outside their confidence bounds.

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

[29] Details on the alternative forecasting model are available in Appendix E.

Figure 3: Ex ante forecasts from the "baseline" model, 2009q1–2011q3

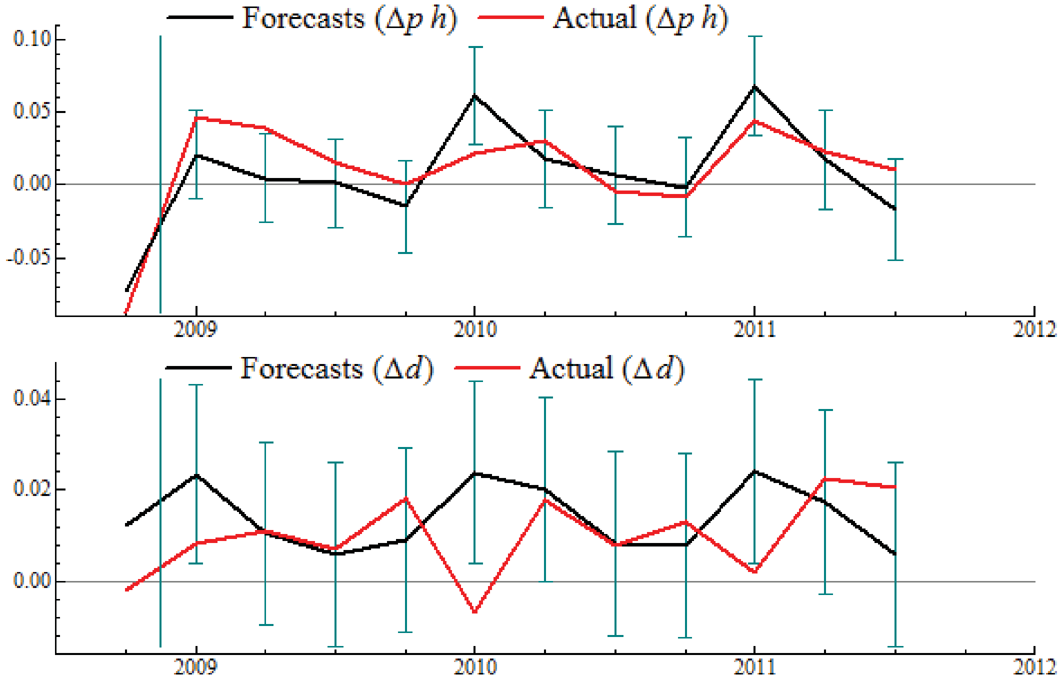
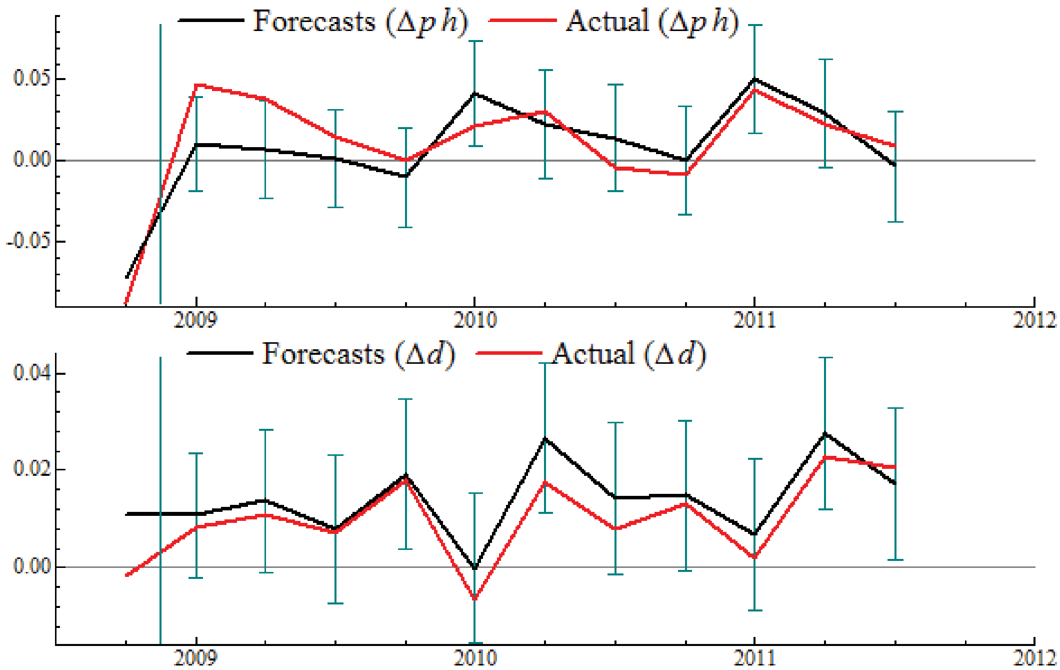


Figure 4: Ex ante forecasts from the model *without* short-run price homogeneity, 2009q1–2011q3



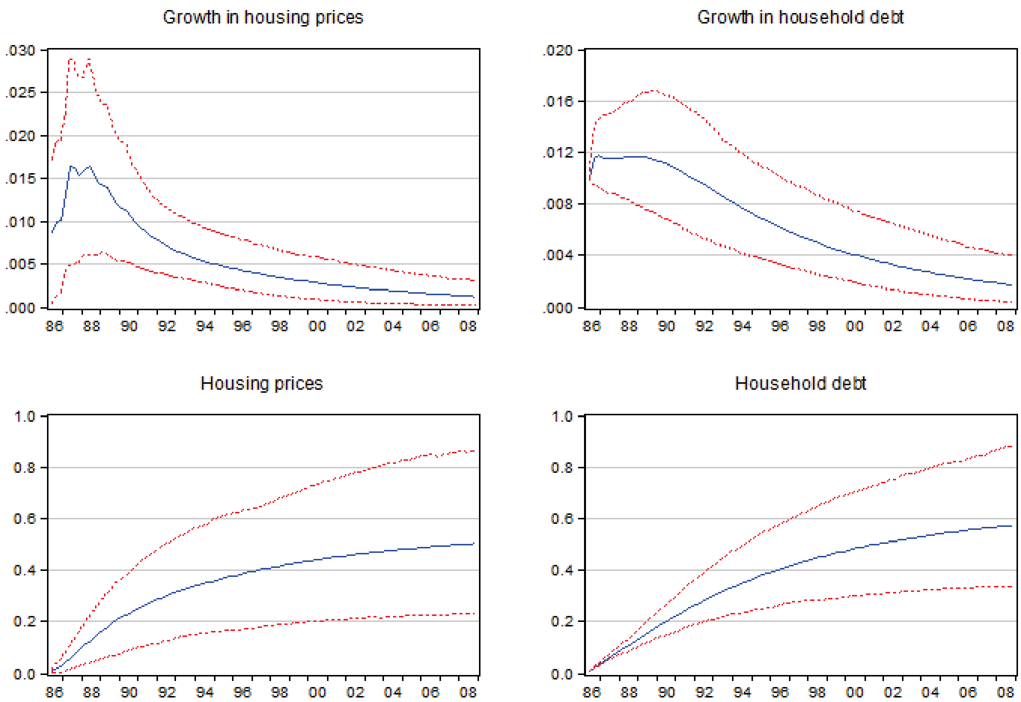
7 Dynamic effects of shocks

In the previous section, we used a general-to-specific approach to specify a parsimonious system capturing the dynamic interaction between housing prices and credit. In the following we will use Monte Carlo simulations of this model to show the dynamic responses to exogenous shocks to the system. As a first step, we consider the subsystem of housing prices and credit developed in Section 7.1, where we condition on the supply side of the housing market. In Section 7.2, we augment the subsystem with a small model for the supply side of the Norwegian housing market. This model is simply taken from an existing model for the Norwegian economy, *i.e.* the Statistics Norway forecasting model KVARTS, see Appendix A for details. As will become evident in the following subsections, including the supply side dampens the long-run impact of shocks, as construction activity responds to changes in housing prices.

7.1 Dynamic multipliers: The baseline model

The first set of simulations we perform are based on the subsystem of housing prices and credit presented in Section 6. All simulations are conducted using 1000 stochastic Monte Carlo replications and 95 percent simulated confidence intervals (dotted red lines) are reported along with the simulated response path (solid blue lines). The dynamic effects of a permanent increase in the growth of credit and housing prices are shown in Figure 5 and 6, respectively. The figures display the impact on the growth rates as well as on the level of real housing prices and the stock of real household debt.

Figure 5: Baseline model dynamic multipliers of a shock to credit growth of 1 percentage point



The figures show that an exogenous shock in one of the markets is propagated and amplified through an endogenous feedback mechanism. Figure 5 shows that a positive exogenous shock in the credit growth by one percentage point will increase housing price

Figure 6: Baseline model dynamic multipliers of a shock to housing price growth of 1 percentage point

growth by 0.86 percentage points at the time of the shock, which equals the instantaneous impact on housing price growth reported in Table 5. The increase in housing prices leads to a further increase in credit growth in the subsequent period, as the collateral value has increased. This again induces further growth in housing prices and credit in a process that continues for about two years before the equilibrium 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, but we see that the levels of both variables have stabilized at a higher level in accordance with the finding of a long-run interaction between housing prices and credit in Section 5. Shocking housing price growth (see Figure 6) yields qualitative effects that are similar to the above described effects, and will of course not change any of the growth rates in the long-run.

A shock to one of the exogenous variables in the system will have similar effects as is shown in Figure 7. A one percent increase in disposable income will lead to a growth in both housing prices and credit, which is reinforced by the feedback between the two variables. The dynamic process 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 the credit market, stimulating housing prices further, and so on. As the cumulative multipliers illustrate, both housing prices and credit continue to grow before the growth rates eventually return to zero. This has of course lead to a new equilibrium price level and a higher fundamental value for the credit variable, as seen from the lower part of the figure. An increase in disposable income, which is one of the long-run determinants of housing prices, will change housing prices and credit period after period until they have

Figure 7: Baseline model dynamic multipliers of an increase in real disposable household income by 1 percent

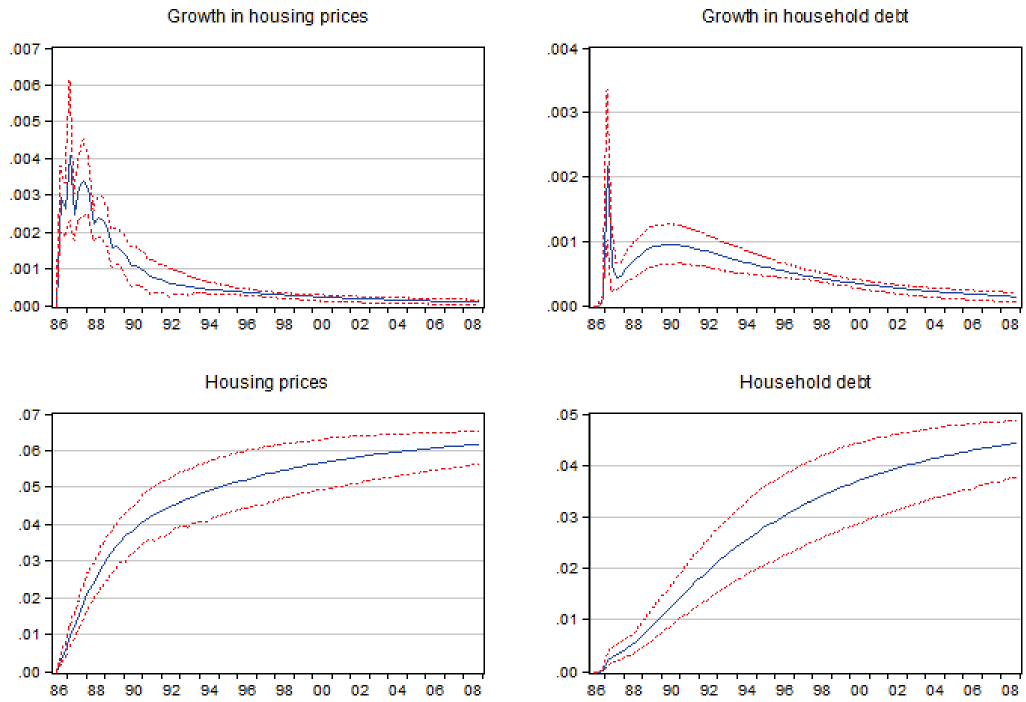
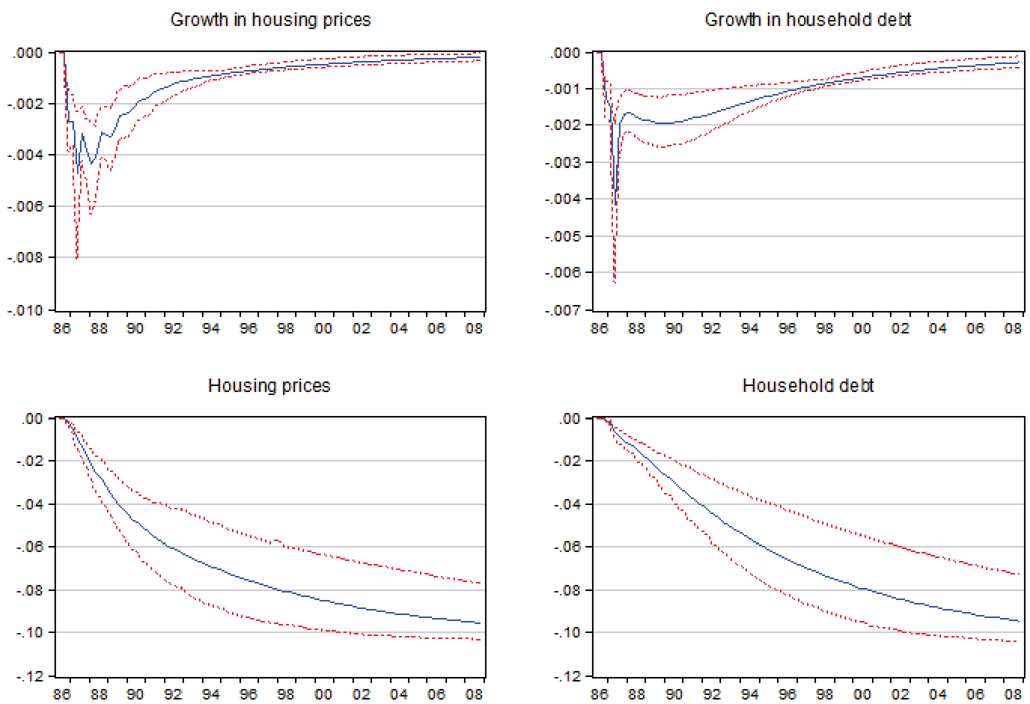


Figure 8: Baseline model dynamic multipliers of a shock to the interest rate of 1 percentage point



adjusted to their new long-run equilibrium level.

Figure 8 shows the simulated responses to one percentage point increase in the real interest rate. This reduces both housing prices and credit growth in the short-run. In the long-run, both housing prices and household debt converge to new and lower equilibrium levels (lower part of the figure), which shows that the model implies interest rates effects on housing prices even though the interest rate does not enter the short nor the long-run equations for housing prices directly.

7.2 Dynamic multipliers: An extended model

In this section we augment the core model above with a small model for the supply side of the housing market. These equations are lifted out of the macroeconometric forecasting model KVARTS, which is an operative and relevant model for the Norwegian economy. The supply side model captures the feedback from housing prices to the investments in new houses, which again affects the housing stock and therefore is expected to dampen the dynamic effects found in the previous subsection. The housing supply model is reestimated on our sample and a brief description of the supply side model, along with the estimated coefficients, are given in Appendix A. Figure 9 and Figure 10 illustrate the dynamic impact of a one percentage point increase in credit growth and housing price growth when the supply side is taken into account.

Though the short-run effects are very similar to those for the baseline model, we see that the effects of the shocks on the growth rates die off more quickly when taking into account that the investment activity responds to changes in housing prices. While in the baseline model a 1 percentage point increase in credit growth still has a great effect on the housing price growth after 4 years, we find that the estimated effect on housing price growth is zero in the extended model after the same period. It follows that also the long-run impact on housing prices and credit is much reduced, as is seen from the graphs in the middle part of Figures 8 and 9. In the long-run, we see the expected convergence to a new equilibrium with higher housing prices and a greater housing stock.

In Figure 11, we have graphed the simulated responses when we increase household disposable income by 1 percent. Again, it is clear that including the supply side dampens the effects relative to those reported in the previous section. In the baseline model, this income shock leads to an increase in housing prices of more than 4 percent after 4 years, and in the long-run the estimated effect on housing prices is around 6%. This contrasts the extended model, where the effect on housing prices after 4 years is around 3%. At this point, the effect gradually declines, as the investment activity increases. In the long-run, we find that housing prices have increased by 0.05 percent, which is half of the initial increase in income. Household debt is found to increase by 1 percent, meaning that the long-run effect on debt will equal the initial shock to income.

The final figure (Figure 12) shows the effect of an increase in the real interest rate of one percentage point. Again, the short-run response is similar to that in the baseline model, while the long-run effect is much reduced. It should be noted that the disposable income variable includes net interest rate income, which is negative on aggregate for the households. Thus, if we had used a larger model, where also disposable income had been modelled, the simulated interest rate effect would be stronger.

Figure 9: Dynamic multipliers of a shock to credit growth of 1 percentage point in the extended model.

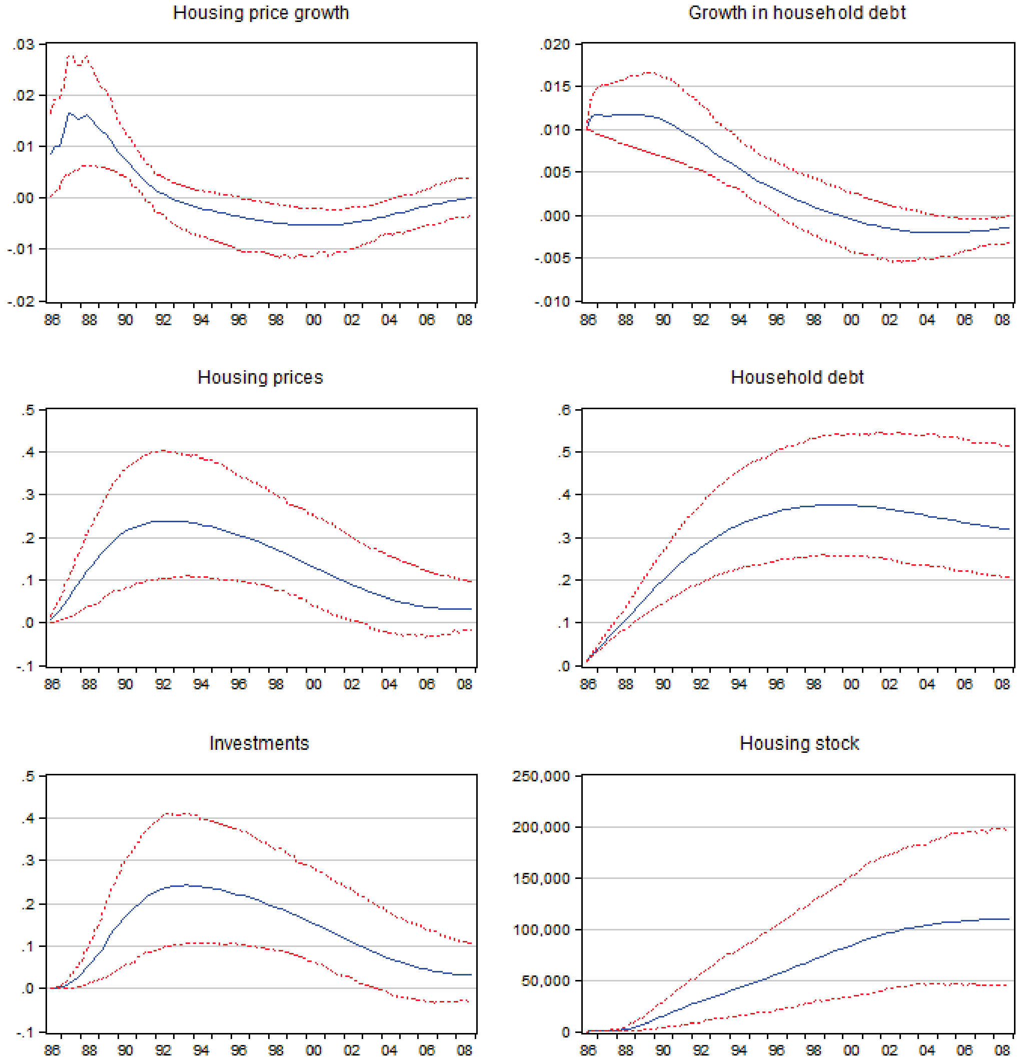


Figure 10: Dynamic multipliers of a shock to housing price growth of 1 percentage point in the extended model.

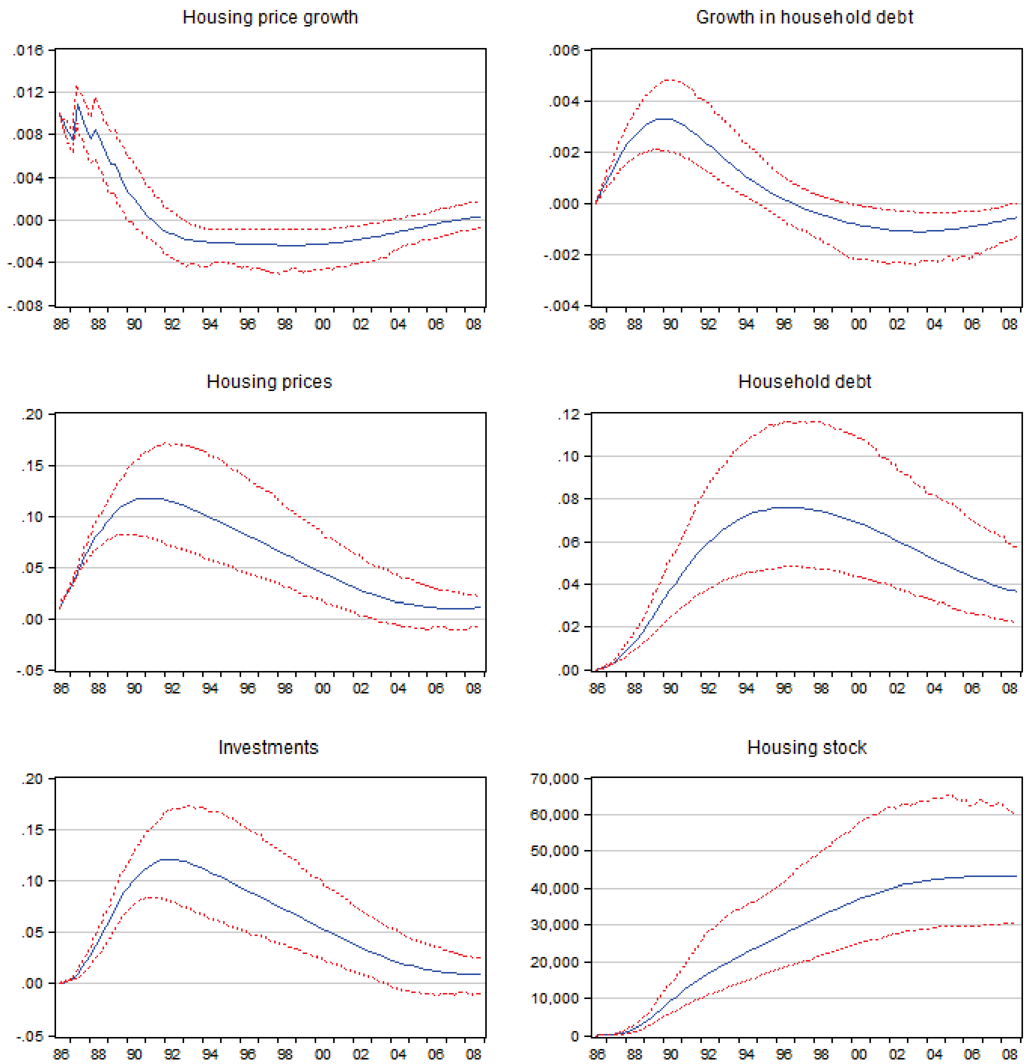


Figure 11: Dynamic multipliers of an increase in real disposable household income of 1 percent in the extended model.

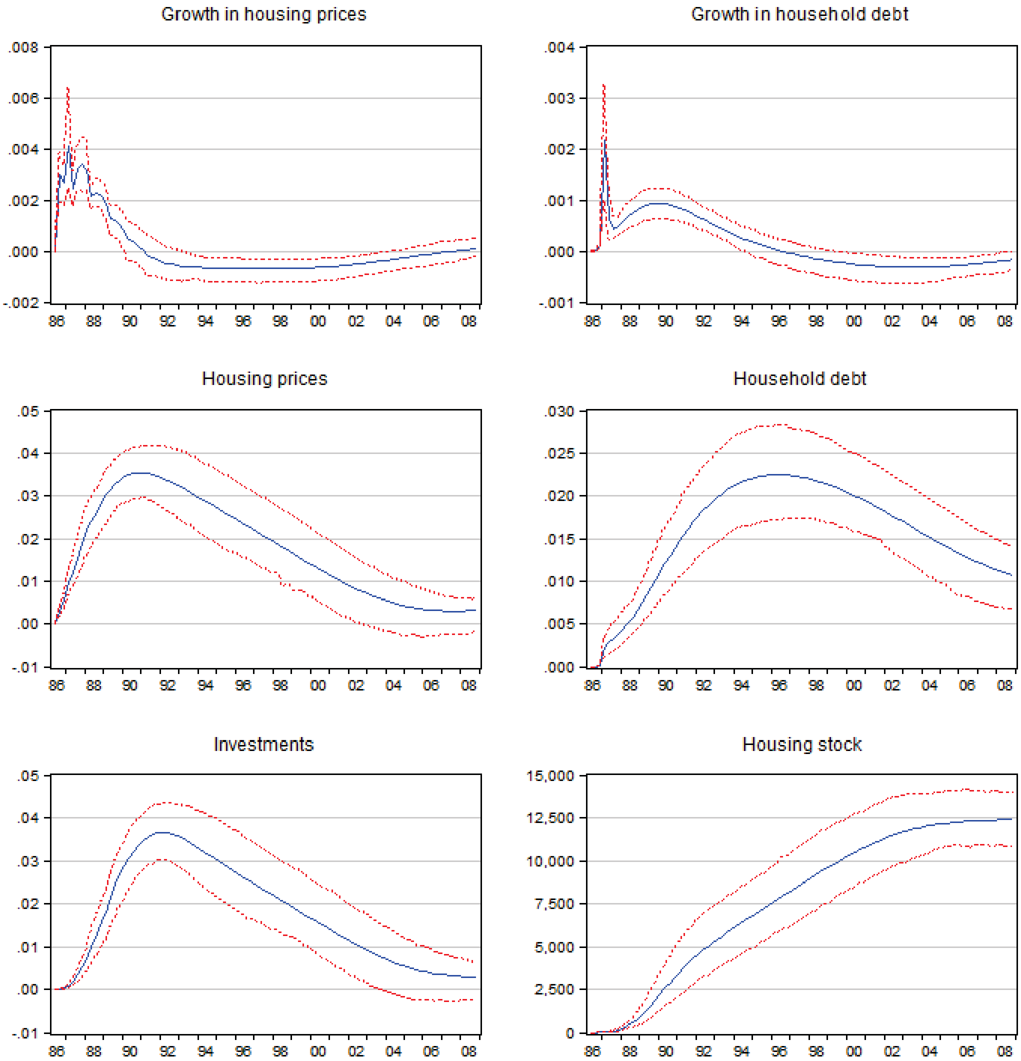
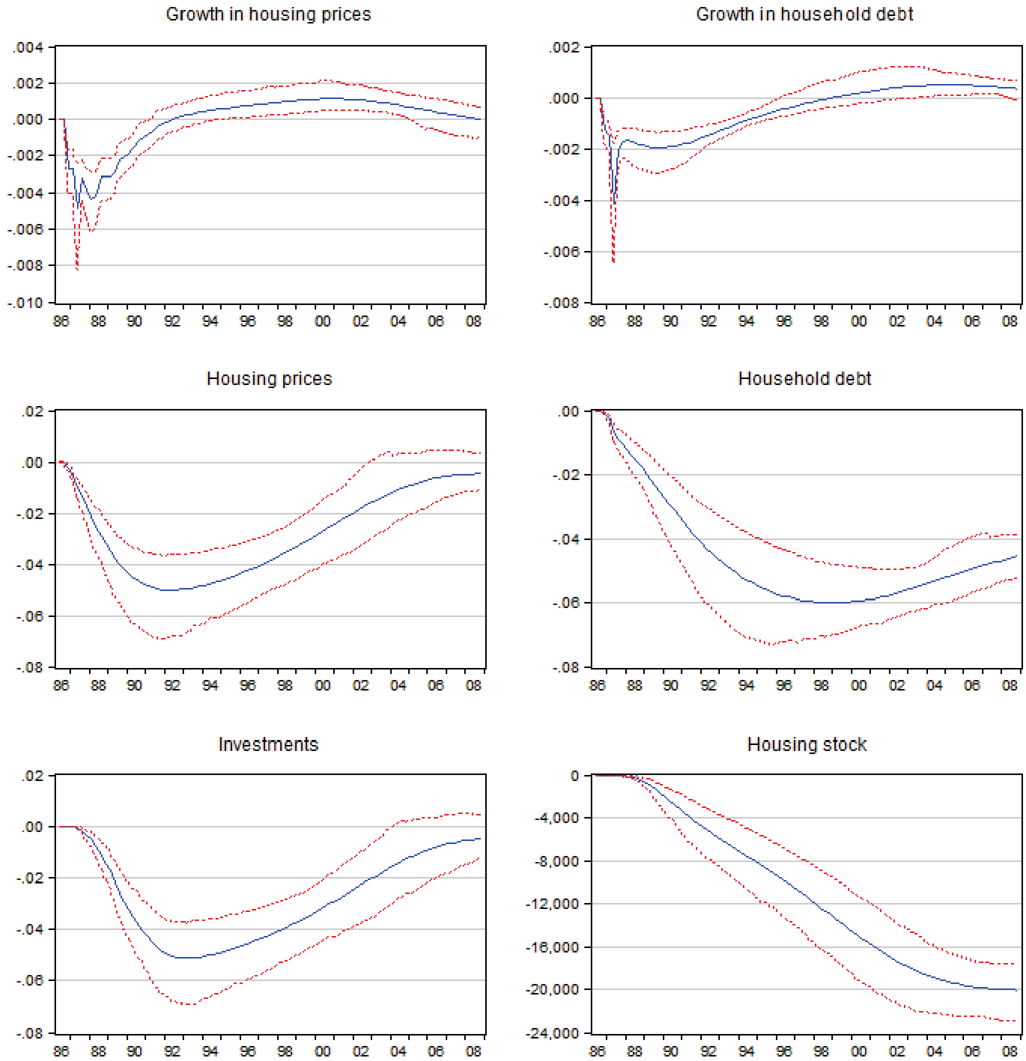


Figure 12: Dynamic multipliers of an increase in the real interest rate of 1 percentage point in the extended model



8 Robustness: Estimating the model on an extended sample

With the benefit of having access to a four more years of data, we have reestimated the short-run dynamics of the Anundsen and Jansen (2013b) model for every quarter between the period 2008q4–2012q4. In addition to having an extended data set, there have also been revisions to the data we originally used. Thus, such a reevaluation of the model is useful to explore the robustness of our results. The recursive coefficient estimates are reported in Figure 13 and 14.

Figure 13: Recursively estimated coefficients for ΔDph equation from the "baseline" model, 2008q4–2012q4

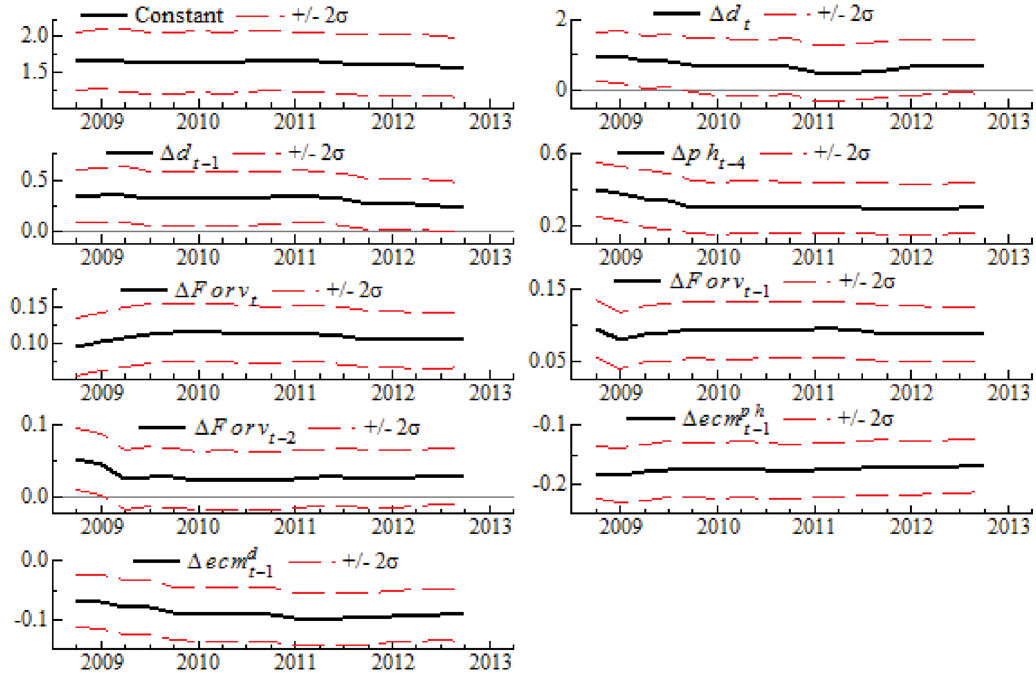
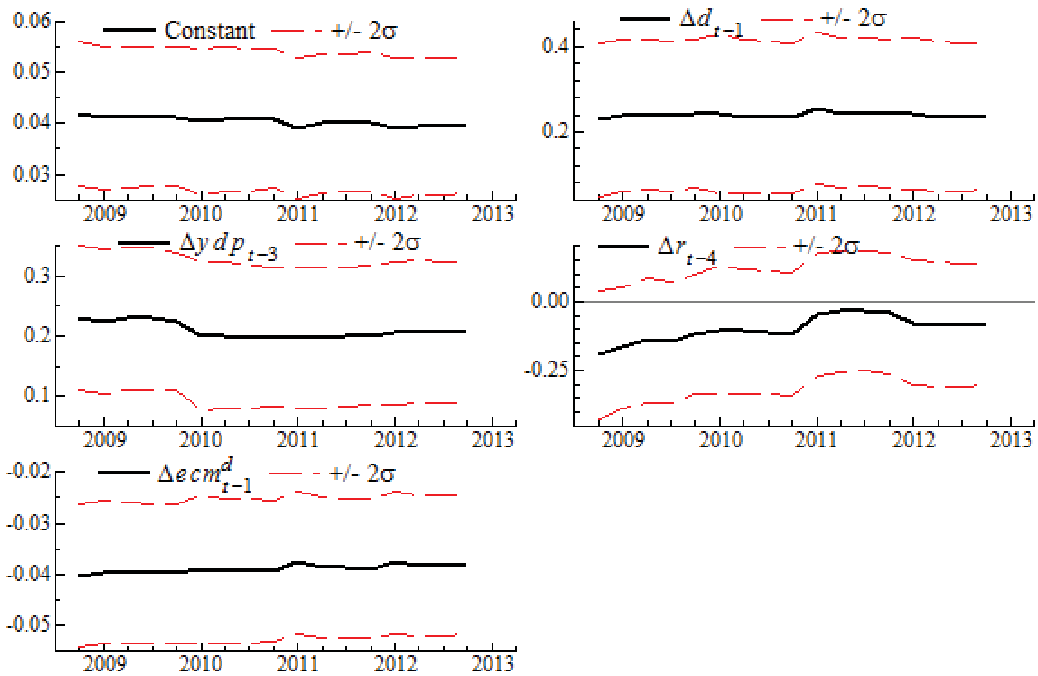


Figure 14: Recursively estimated coefficients for ΔDd equation from the "baseline" model, 2008q4–2012q4



It is clear that all the coefficients are stable when estimated recursively. The same recursive estimates for the model *without* short run price homogeneity are reported in Figure E.1 and E.2 in Appendix E and they show the same picture. This is a reassuring finding, and is particularly important if the model is to be used for forecasting purposes. Having a good forecasting model for housing prices and credit seems imperative both in order to monitor the development in the financial system and to increase the forecasting accuracy of key macroeconomic variables such as consumption and investments. In fact, preliminary results (Anundsen and Jansen, 2013a) show that the *ex ante* forecasts from the model documented in this chapter fares well against alternative forecasting models, such as autoregressive, vector autoregressive and random walk models.

Finally, again with the benefit of having access to more data, Figure 15 and 16 show the forecasts for the model with and the model without short run price homogeneity from 2008q4–2012q4, i.e. adding five more observations to the forecasting horizon relative to what we did in Section 6

Figure 15: Ex ante forecasts from the "baseline" model, 2009q1–2012q4

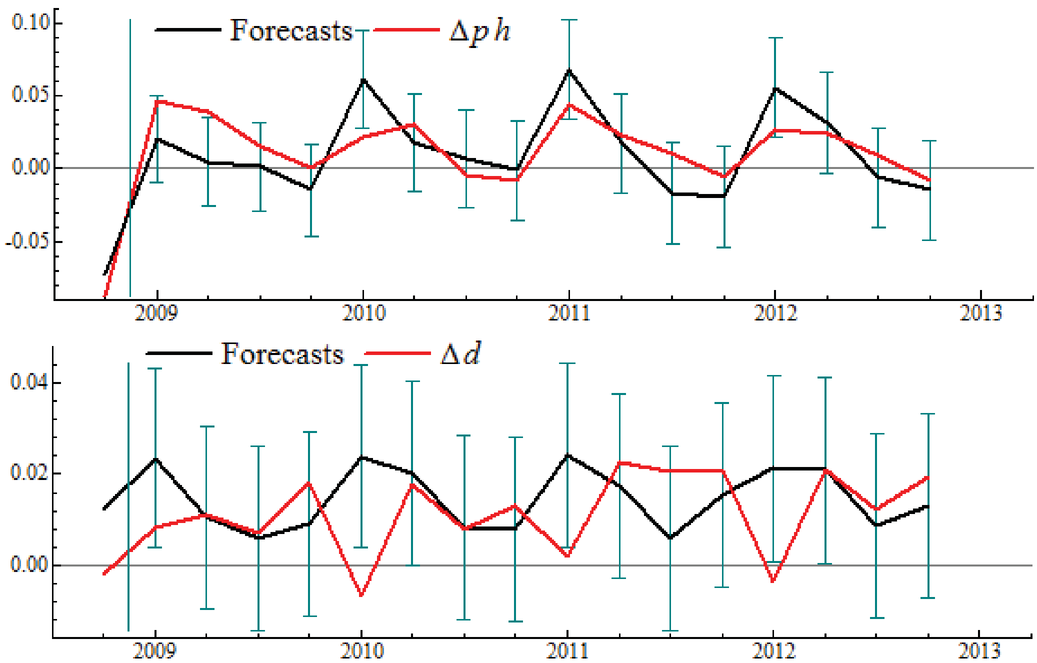
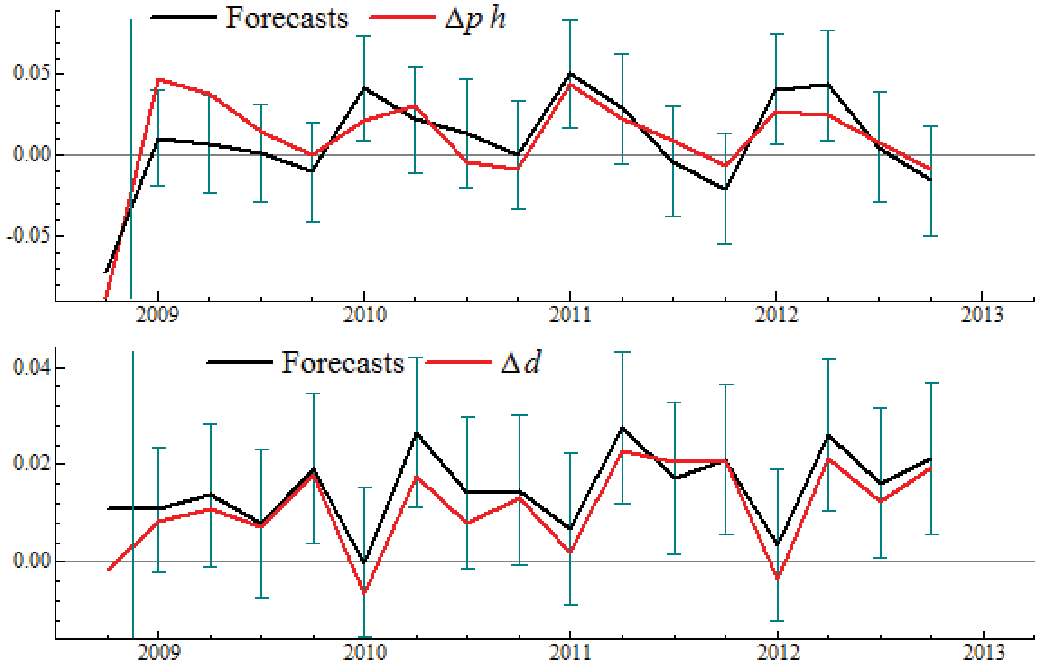


Figure 16: Ex ante forecasts from the model *without* short-run price homogeneity, 2009q1–2012q4



Also for the last year, we see that the credit forecasts produced by the baseline model are outside their confidence bounds in the first quarter. However, the model where we have derestricted the short-run price homogeneity does a far better job, which underpins our argument that including short-run inflation effects in the credit equation may be important for forecasting purposes. In conclusion, it seems that the model passes the stability tests when evaluated on an extended sample.

9 Conclusion

Using cointegration analysis, this study documents the importance of jointly estimating long-run interactions between house prices and household debt. Furthermore, estimating these variables in a vector error-correction system also yields better estimates of short-run interactions and dynamic responses. We find evidence that household income is weakly exogenous with respect to other long-run housing-related variables. Along with other tested constraints on coefficients, we use this finding to estimate a more parsimonious system of household debt and house prices.

In particular, we find that house 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 disequilibria in the credit market.

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

Third, a consumer confidence indicator measuring households' expectations concerning future developments in their own economy as well as the Norwegian macro economy are 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 analysis of the dynamic multipliers 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. Incorporating a model of the supply side of the housing market dampens the dynamic responses of housing prices and credit to all shocks considered here. This highlights the importance of accounting for construction, as well as credit factors, in modeling housing cycles.

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Appendix A: The supply side

The equations describing the supply side of the housing market in Section 7 are lifted out of the Statistics Norway quarterly forecasting model, KVARTS (Eika and Moum (2005)) and reestimated on the current sample (1986q2-2008q4). In KVARTS, the supply of housing is modelled by considering housing starts measured in square meters. Housing starts serve as a leading indicator for the development in housing investments, which eventually become new houses and add to the housing stock.

In a long-run perspective, new housing starts are modeled according to the q-theory of investments, where a one percent increase in either housing prices or a one percent decrease in construction costs lead to a one percent increase in housing starts. This implies that a proportional increase in construction costs and housing prices will have no long-run effect on the supply of new houses. Letting S denote housing starts, PJ denote real construction costs and PH denote real housing prices, the reestimated equation for housing starts is given by (absolute t-values reported under the point estimates).

$$\begin{aligned} \Delta logS_t &= \underset{(4.90)}{0.41} \Delta logS_{t-4} - \underset{(4.28)}{0.26} \left(logS_{t-1} - logPH_{t-1} + logPJ_{t-1}\right) \\ &\quad + dummies \\ R^2 &= 0.77 \end{aligned} \tag{A.1}$$

In addition to the equilibrium correction term, the model contains an autoregressive part as well as an impulse dummy for the second quarter of 2002 and a set of seasonal dummies for the first three quarters. The re-estimated coefficients are almost unchanged from the version used in KVARTS, which is reassuring.

Since it takes time for a newly started building project to get finished, it is assumed that a change in housing starts will lead to a flow of investments for several years. In KVARTS this adjustment is assumed to take 12 quarters and the relationship linking investments and housing starts is given by the following equation:

$$\Delta log(IH) = \Delta log(J) + seasonals \tag{A.2}$$

where IH denotes housing investments, which grow proportionally with a weighted average of housing starts over the last 12 quarters, J. Also the coefficients for the seasonal dummy variables in equation (A.2) are reestimated when we construct the model used for simulations in Section 7. The weighted average of housing starts is given by the following identity .

$$\begin{aligned} J = {} & 0.3124 * S_t + 0.2455 * S_{t-1} + 0.1672 * S_{t-2} + 0.1125 * S_{t-3} + 0.0702 * S_{t-4} \\ & + 0.0407 * S_{t-5} + 0.0235 * S_{t-6} + 0.0131 * S_{t-7} + 0.0074 * S_{t-8} + 0.0043 * S_{t-9} \\ & + 0.0021 * S_{t-10} + 0.009 * S_{t-11} + 0.002 * S_{t-12} \end{aligned}$$

Finally, the housing stock is determined by a law of motion of capital accumulation:

$$H_t = (1 - \delta)H_{t-1} + IH_t$$

where δ is the rate of depreciation of the housing stock. As is evident from this brief presentation of the supply side, the model used for simulation in Section 7 captures spill overs from housing prices to the construction sector, which, as shown in the simulation exercises, dampens the long-run effect of shocks on housing prices and credit.

Appendix B: 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 interest rate paid by households on loans in private financial institutions. Source: Statistics Norway.
p: Consumer Price Index. Source: Statistics Norway.
π*:* Annual inflation rate ($\Delta_4 p$).
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 ($i * (1 - \tau) - \pi$).

Variables used in Appendix A:
S: Housing starts (square meters). Source: Statistics Norway.
J: Weighted sum of housing starts (square meters).
IH: Investments in housing, measured at fixed prices. Source: Statistics Norway.
PJ: Price index for construction costs, deflated by *pc*. Source: Statistics Norway.
δ*:* rate of depreciation of the housing stock.

Appendix C: Tables

Table C.1: Tests for the order of integration[a]

	ADF		PP		KPSS		
Testing levels							
Variable	$t - ADF$	5%	$Adj.t - stat$	5%	LM	5%	Characteristics[b]
ph	−2.37	−3.46	−1.32	−3.46	0.27	0.146	t
d	−3.77	−3.46	−0.69	−3.46	0.27	0.146	t
h	−2.76	−3.46	−0.78	−3.46	0.22	0.146	t
yh	−0.98	−3.46	−5.18	−3.46	0.31	0.146	t
th[c]	−3.21	−3.46	−7.74	−3.46	0.14	0.146	t
r	−3.58	−3.46	−3.5	−3.46	0.13	0.146	t
E[d]	−1.80	−3.46	−2.15	−3.46	0.08	0.146	t
Testing first differences							
Δph	−2.07	−2.89	−5.99	−2.89	0.25	0.46	i
Δd	−1.77	−2.89	−5.35	−2.89	0.3	0.46	i
Δh	−2.198	−2.89	−1.84	−2.89	0.29	0.46	i
Δyh	−4.25	−2.89	−27.05	−2.89	0.44	0.46	i
Δth	−8.71	−2.89	−21.91	−2.89	0.11	0.46	i
Δr	−11.11	−2.89	−10.73	−2.89	.10	0.46	i
ΔE	−5.12	−2.89	−7.55	−2.89	0.28	0.46	i
Testing second differences							
$\Delta^2 ph$	−4.62	−2.89	–	–	–	–	i
$\Delta^2 d$	−13.28	−2.89	–	–	–	–	i
$\Delta^2 h$	−2.548	−2.89	−11.41	−2.89	–	–	i

[a] While the PP and KPSS tests are performed in EViews, we run the ADF test in PcGive since this allow us to include seasonal dummies in the test regression. The variables for which we have included seasonal dummies in the test regressions are housing prices, disposable income and the turnover, as they all display a clear seasonal pattern. When inspecting this table, it is important to keep in mind that while the ADF test and the PP test have non-stationarity as the null, the KPSS test has stationarity as the null.

[b] The different characteristics are: Including both trend and intercept (t) or only an intercept (i) in the test regression.

[c] The turnover is only collected from 1985q1, which means that with 8 lags in the ADF regression, the sample starts in 1987q2.

[d] For the expectations variable we only have data for the period from 1992q3 and the variable is set to 0 in the period prior to this in the empirical analysis. For the tests for the order of integration, we use the period for which we have observations.

Table C.2: 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.

Table C.3: Loading factors for the models reported in Table 4

Loading	Panel 1	Panel 2	Panel 3	Panel 4	Panel 5
$\alpha_{1,ph}$	-0.82	-0.76	-0.21	-0.22	-0.24
	0.21	0.20	0.04	0.04	0.04
$\alpha_{1,d}$	-0.44	-0.40	-0.06	-0.08	-0.1
	0.14	0.13	0.02	0.03	0.03
$\alpha_{2,ph}$	-0.13	-0.13	0	0	0
	0.11	0.11	$--$	$--$	$--$
$\alpha_{2,d}$	-0.11	-0.10	-0.04	-0.05	-0.04
	0.07	0.07	0.01	0.01	0.01
$\alpha_{3,ph}$	0.42	0.40	-0.04	-0.05	0
	0.15	0.14	0.03	0.03	$--$
$\alpha_{3,d}$	0.31	0.29	0.002	-0.01	0
	0.1	0.09	0.02	0.02	$--$

Note: This table reports the estimated loading factors (equilibrium correction coefficients) obtained when we impose the various overidentifying restrictions on our two cointegrating vectors, confer Table 4 for the estimated cointegrating vectors. The second rows contain for each loading the estimated standard errors across panels.

Table C.4: Augmented Dickey-Fueller tests for structural residuals[a]

Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
$\varepsilon_{\Delta ph}$	-8.846	-2.89	0	No	No
$\varepsilon_{\Delta d}$	-7.945	-2.89	1	No	No

[a] The residuals from the short run system is tested over the period 1988q3-2008q4 since we only obtain data for the error correction terms from 1986q2.

Appendix D: Equation-by-equation modelling

Adopting a single equation approach one would take the system represented by equation (9) and (10) as a starting point. This approach precludes any formal treatment of identification, but may possibly give reasonable results if the simultaneity bias is not large. We have used the automated multipath search algorithm Autometrics (see Doornik (2009) and Doornik and Hendry (2009b)) to reduce the dimensionality of each equation. An obvious advantage with this algorithm is that it is very little path dependent as it does a multipath search. However, the benefit from this might be outweighed by the fact that it does not allow us to take care of the simultaneity from the onset by doing a full fledged system analysis at each step in the reduction process. The results from this single equation general to specific approach are documented in Table D.1 and Table D.2 for the housing price and credit equation, respectively.

Table D.1: Short run dynamics obtained by Autometrics for housing price equation[a]

Variable	Coefficient	t-value
$Constant$	1.23	6.78
Δd	0.61	3.85
Δph_{t-4}	0.41	4.93
Δt_{t-3}	0.05	2.55
Δr_{t-4}	-0.38	2.06
ΔE_t	0.095	4.54
ΔE_{t-1}	0.096	4.40
ΔE_{t-2}	0.05	2.17
ecm^{ph}_{t-1}	-0.07	3.81
ecm^{d}_{t-1}	-0.14	6.80
$CSeasonal_t$	-0.006	0.496
$CSeasonal_{t-1}$	-0.007	0.65
$CSeasonal_{t-2}$	-0.009	0.999
σ	0.0141	
R^2	0.82	
$Adj.R^2$	0.80	
Diagnostics[b]	Test statistic	Value [p-value]
AR 1-5 test:	$F(5,73) =$	0.4789 [0.7909]
ARCH 1-4 test:	$F(4,83) =$	0.4462 [0.7749]
Normality test:	$\chi^2(2) =$	1.5603 [0.4583]
Hetero test:	$F(21,69) =$	1.3658 [0.1672]
Estimation Method	OLS (Autometrics with p-value = 0.05)	
Sample	1986q2-2008q4	

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

Table D.2: Short run dynamics obtained from Autometrics for the credit equation[a]

Variable	Coefficient	t-value
$Constant$	-0.73	10.6
Δph_t	0.30	7.06
Δph_{t-4}	-0.12	2.64
Δy_{t-2}	-0.15	3.10
ΔE_{t-1}	-0.04	2.45
Δr_{t-3}	-0.24	2.34
ecm^{ph}_{t-1}	0.09	10.8
$CSeasonal_t$	-0.004	1.16
$CSeasonal_{t-1}$	-0.004	1.50
$CSeasonal_{t-2}$	-0.01	4.07
σ	0.009	
R^2	0.72	
$Adj.R^2$	0.69	
Diagnostics[b]	Test statistic	Value [p-value]
AR 1-5 test:	$F(5,76) =$	1.4959 [0.2011]
ARCH 1-4 test:	$F(4,83) =$	0.7501 [0.5608]
Normality test:	$\chi^2(2) =$	4.9864 [0.0826]
Hetero test:	$F(15,75) =$	0.8092 [0.6641]
Estimation Method	OLS (Autometrics with p-value = 0.05)	
Sample	1986q2-2008q4	

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

The results in Table D.1 and Table D.2 reveal some differences as compared to our preferred model. We note that both variables enter contemporaneously in both equations. Also, we observe that the income variable and the expectations variable are both highly significant in the credit equation with negative signs, which are not plausible *a priori* . Let us now turn to the two equations when they are estimated simultaneously to take care of potential endogeneity problems. Results are displayed in Table D.3.

Table D.3: System estimation of the specifications obtained by Autometrics (equation by equation)[a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.00	3.78	−0.73	10.5
Δd_t	−0.26	0.49	–	–
Δph_t	–	–	0.32	5.50
Δph_{t-4}	0.36	3.65	−0.13	2.57
Δyh_{t-2}	–	–	−0.15	3.05
ΔE_t	0.12	3.88	–	–
ΔE_{t-1}	0.10	3.95	−0.04	2.48
ΔE_{t-2}	0.05	1.75	–	–
Δr_{t-3}	–	–	−0.24	2.37
Δr_{t-4}	−0.51	2.36	–	
Δt_{t-3}	0.06	2.50	–	
ECM^{ph}_{t-1}	−0.11	3.34	0.09	10.6
ECM^{d}_{t-1}	−0.10	3.85	–	–
Dummy, q1	−0.01	0.75	−0.005	1.26
Dummy, q2	−0.009	0.73	−0.004	1.55
Dummy, q3	−0.02	1.61	−0.01	4.07
Sargan	$\chi^2(43) =$	40.323 [0.5881]		
Log likelihood	567.99			
σ	0.016		0.0086	
Diagnostics[b]	Test statistic	Value [p-value]		
Vector SEM-AR 1-5 test:	$F(20,138) =$	0.7944[0.7168]		
Vector Normality test:	$\chi^2(4) =$	4.7544[0.3134]		
Vector Hetero test:	$F(183,81) =$	1.0260[0.4557]		
Estimation Method	FIML			
Sample	1986q2-2008q4			

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

The credit equation remains almost unaltered, while the housing price equation changes dramatically. First of all, the credit variable which is positive and highly significant in the single equation model has now changed sign and is insignificant. Also, the loadings have changed. As a final check of this model, we will explore how the implied dynamics of the system to a permanent increase in real disposable income would be. We follow exactly the same set up as in section 7.1 of the paper and the dynamic multipliers are graphed in Figure D.1.

Figure D.1: The alternative model: Dynamic multipliers of a 1 percent increase in real disposable household income.

Based on the dynamic multipliers from this alternative model, we see that it implies a negative response to household borrowing of an increase in income in the short run, which seems unreasonable from an economic point of view. Also, the credit effect on housing prices changes sign and turns out insignificant, though it was positive and highly significant in the single equation case. Furthermore, we observe relative big changes in the loadings in the housing price equation. On this background we conclude that this model is inferior to the one from the simultaneous model design reported in Table 5 in Section 6 of the paper.

Appendix E: Model without short run price homogeneity

With reference to the forecasting exercise in Section 6 of the paper, this section discusses a version of the model, where we de-restrict the assumption of short run price homogeneity. To see whether the forecast failures for the credit growth in 2010q1 and 2011q1 (confer Figure 2 in the paper) may be due to the extremely cold winters, which lead to an extraordinary jump in electricity prices in each of the two quarters, we re-estimated the model for the case where short run price homogeneity is relaxed. As shown in the paper (see Figure 3), this improves the forecasting accuracy of the model – and in particular the credit forecasts. The estimation results underlying those forecasts are reported in Table E.1.

Table E.1: Short run dynamics from the model *without* short-run price homogeneity[a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.617	7.90	0.023	4.83
Δd_t	0.696	3.78	-	-
Δd_{t-1}	-	-	0.560	7.68
Δd_{t-3}	0.355	2.69	-	-
Δph_{t-4}	0.394	5.07	-	-
Δyh_{t-3}	-	-	0.084	1.99
ΔE_t	0.102	5.12	-	-
ΔE_{t-1}	0.100	4.76	-	-
ΔE_{t-2}	0.045	2.05	-	-
ΔR_{t-4}	-	-	-0.088	1.13
Δpc_t	-	-	-0.720	9.25
Δpc_{t-1}	-	-	0.528	5.89
ECM^{ph}_{t-1}	-0.172	7.86	-	-
ECM^{d}_{t-1}	-0.071	4.26	-0.025	4.63
Dummy, q1	0.025	3.87	-0.016	4.37
Dummy, q2	0.024	4.27	0.007	2.52
Dummy, q3	0.013	2.31	-0.019	7.36
Sargan		$\chi^2(48) =$	44.68 [0.6099]	
Log likelihood		603.68		
σ	0.0137		0.0064	
Diagnostics[b]		Test statistic	Value [p-value]	
Vector EGE-AR 1-5 test:		F(20,138)	0.50 [0.96]	
Vector Normality test:		$\chi^2(4)$	36.17 [0.00]	
Vector hetero test:		F(195,69)	0.67 [0.98]	
Estimation Method	FIML			
Sample	1986q2-2008q4 ($T = 91$)			

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

We started by including the current and first lag of the change in the price deflator (Δpc) in both equations. However, these variables were only significant in the credit equation, and were therefore excluded from the housing price equation. As seen, the inclusion of Δpc_t and Δpc_{t-1} in the credit equation only has minor effects on the estimated parameters of the housing price equation, while the estimates of the credit equation are somewhat changed. That said, it seems to be changed for the better, since – as is evident from inspecting the table – derestricting short run price homogeneity improves the fit of the credit equation. Furthermore, both the current and lagged value are highly significant, and come with opposite signs. In fact, we can not reject the hypothesis that the two coefficients are equal in absolute value, i.e. suggesting that these terms are measuring a surprise inflation ($\Delta^2 pc_t = \Delta pc_t - \Delta pc_{t-1}$). This gives additional credence to our conjecture that the forecast failures in 2010q1 and 2011q1 are due to an unexpected increase in electricity prices.

Recursive estimates for this model for the period 2008q4–2012q4 are displayed in Figure E.1 and E.2.

Figure E.1: Recursively estimated coefficients for Δph equation from the model *without* short-run price homogeneity, 2008q4–2012q4

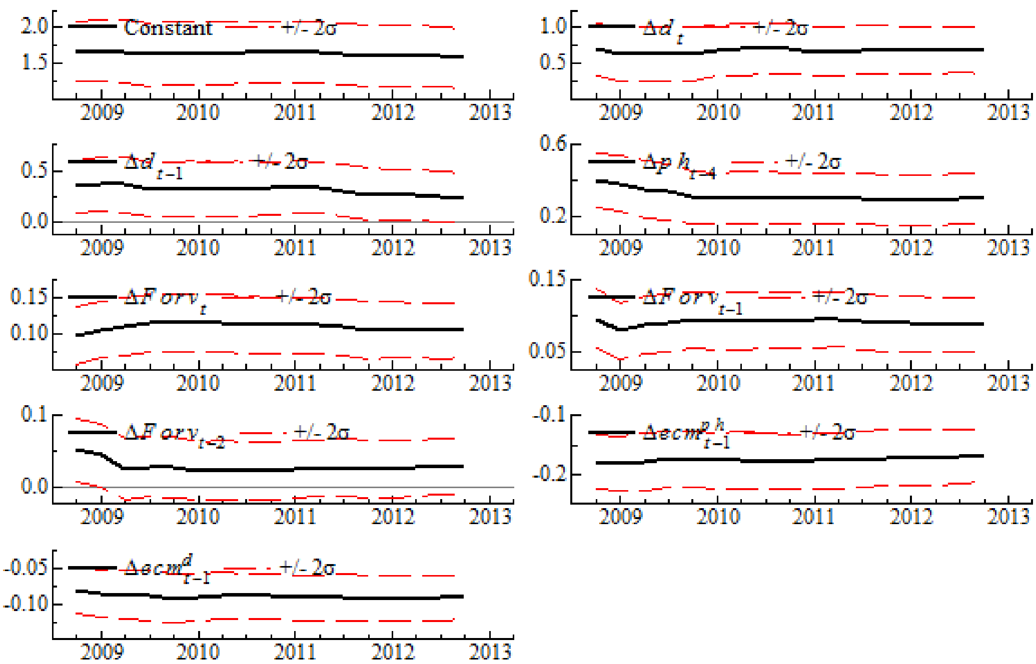
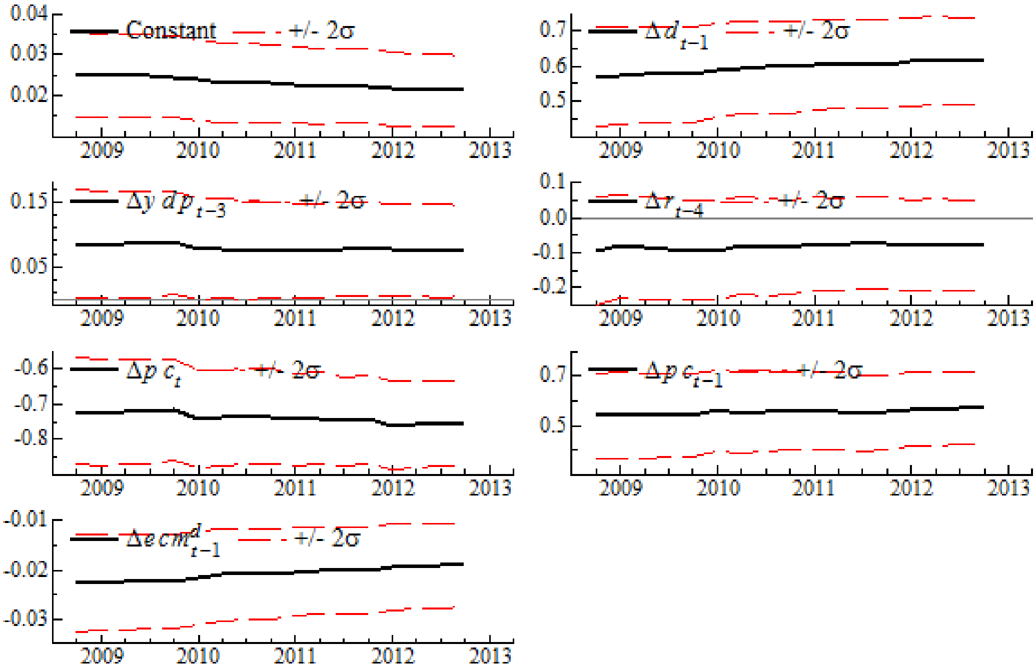


Figure E.2: Recursively estimated coefficients for Δd equation from the model *without* short-run price homogeneity, 2008q4–2012q4



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