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ABSTRACT

This article provides new empirical evidence on the dynamics of price movements and transaction volume in the housing market using data from Finland. While the previous related literature studies the reactions of sales volume and prices to an interest rate shock only, we investigate the responses to income and credit shocks as well. Based on an estimated vector-error correction model, the response of prices to demand shocks is found to be substantially slower than that of sales. The effect of a demand shock on sales peaks within a quarter from the shock. The results show that the differences in the reaction patterns to demand shocks can create the kind of strong positive co-movement between price movements and sales and the kind of negative correlation between price level and sales that has been found in a number of housing markets. It is also found that the direct predictive ability of the fundamentals with respect to housing price growth and sales is overwhelmed by the predictive power of the lagged observations on price changes and sales themselves.

JEL Classification: R21; R31

Keywords: Housing, dynamics, sales, transactions, prices

Contact information

Elias Oikarinen, Department of Economics, Turku School of Economics, Rehtorinpellonkatu 3, 20500 Turku, Finland. Email: elias.oikarinen(at)tse.fi

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1 Introduction

In an efficient and frictionless asset market transaction volume and price movements should be uncorrelated. Prices should adjust to shocks immediately without any change in sales volume or selling times. Nevertheless, there is a vast empirical literature showing significant positive correlation between sales and price movements in the housing market. Similarly, a number of studies have found an inverse relationship between average sales times and price movements of housing. More recently, a lead-lag relationship between sales and price movements has also been found in some housing markets.

On the theoretical side, a number of reasons for the empirical observations have been suggested. Stein (1995) and Genesove and Mayer (1997) relate the co-movement between the sales volume and price movements to households' liquidity constraints. Wheaton (1990) and Krainer (2001), instead, derive search-theoretic models that emphasize the role of market frictions other than liquidity constraints, such as costly search and informational assymetries, to explain the observed phenomena. While Wheaton's and Krainer's models are able to explain the contemporaneous correlation between the variables, Berkovec and Goodman (1996) present a model that yields a lead-lag relatioship between transaction volume and price growth. In addition, the findings by Genesove and Mayer (2001) and Einiö et al. (2008) suggest that irrational loss aversion may play a significant role in the relationship between sales and prices in the housing market, especially in times of declining housing demand. Case and Shiller (2003), in turn, claim that selling times fall in hot markets because overexuberant buyers have irrational house price anticipations, formed by backward-looking expectations.

Importantly, the theoretical considerations and empirical findings of Berkovec and Goodman (1996) propose that data on transaction volume can be used to predict future housing price movements. On the other hand, because of the typically backward-looking expectations in the housing market, housing price movements might be used to predict future evolution of transaction volume. Nevertheless, empirical research on the dynamics between the variables is still limited, and the predictive power of price movements with respect to sales and vice versa has not been rigorously investigated previously. Hence, further research is needed.

This article provides new empirical evidence on the co-movement between and dynamics of prices and sales volume in the housing market using quarterly data over 1975-2008 from Finland. The study contributes to the literature in several ways. First, we estimate vector-error correction models to study the dynamics rigorously, whereas previous related studies have not investigated the dynamics in a vector-error correction framework. Second, a particular emphasis is given to examining the reaction speeds of prices and sales to shocks in the fundamental variables affecting housing demand. While the previous related literature studies the reactions of prices and sales to an interest rate shock only (see Hort, 2000; Andrew and Meen, 2003a), this article investigates the reactions to income and credit shocks as well. Furthermore, the out-of-sample forecasting properties of housing price movements and sales volume with respect to each other are examined.

The estimated reaction speed of sales to a demand shock is somewhat different from those presented by Hort (2000) and Andrew and Meen (2003a) regarding the Swedish and UK markets, respectively. In particular, the maximum effect of a demand shock on sales takes place within a quarter from the shock, i.e. sales react faster than based on the Hort and Andrew-Meen models. In line with the predictions of the Berkovec-Goodman (1996) model, the response of prices to demand shocks is found to be substantially slower than that of sales. The impulse response analysis also shows that the differences between the reaction patterns and speeds of housing prices and sales to demand shocks can create the kind of strong positive co-movement between price movements and sales and the kind of negative correlation between price level and sales that has been found in a number of housing markets.

While price movements and sales are found to Granger cause themselves and sales also Granger cause price growth, only weak support for positive feedback from price movements to sales is found. Importantly, the predictive power of the lagged observations of price changes and sales overwhelms any direct predictive ability of the lagged changes in the fundamentals that drive housing demand.

The remainder of the paper is organized as follows. The next section reviews previous literature on the co-movement and dynamics between prices and sales in the housing market. Then, the empirical model used in the study is presented. The fourth part outlines the data employed in the empirical analysis, after which the econometric methodology is described. Section six reports the results from the econometric analysis. In the end, the paper is summarized and conclusions are derived.

2 Previous literature

In an efficient and frictionless market, price movements and transaction volume should be uncorrelated. Demand shock, for instance, would induce an immediate change in the price level with no effect on the transaction volume (Andrew and Meen, 2003a). Nevertheless, literature presents several reasons to expect a strong positive relationship between price movements and transaction volume in the housing market. Therefore, it is not unexpected that empirical research has generally found positive co-movement between housing price growth and transaction volume.

Previous empirical literature has studied the interaction between housing prices, housing sales and average selling times mainly concerning the US and UK markets and to some extent also regarding the Swedish market. Among the several articles reporting positive contemporaneous correlation between price movements and the level of sales in the housing market are Stein (1995) Andrew and Meen (2003a), and Ortalo-Magné and Rady (2006). Berkovec and Goodman (1996), in turn, find a positive correlation between price growth and changes in the sales volume. Since there is a strict inverse relationship between transactions and average selling times (average selling time = vacancies / transactions, see e.g. DiPasquale and Wheaton, 1996; Meen, 2001), it is not unexpected that DiPasquale and Wheaton (1996) show evidence for significant negative correlation between sales time and real price growth in the US. Follain and Velz (1995) and Hort (2000), instead, report negative correlation between sales and price level in the US and Sweden, respectively.[1] The specification (levels vs. differences) of the data series employed in the correlation analysis is likely to affect the findings. Furthermore, in the study by Follain and Velz the shortness of the sample period (1986-1993) may have biased the results. Moreover, structural breaks regarding the relationship between sales and prices may occur. Andrew and Meen (2003b) argue that a

[1] Hort finds the contemporaneous co-movement between sales and prices to be insignificant, though.

structural break, which was caused partly by the changed behavior of young individuals, in the relationship between housing prices and transactions in the UK took place during the 1990s.

Given the high transaction costs, time consuming transaction process, relative illiquidity and informational asymmetries in the housing market and the liquidity constraints faced by households, the assumption of frictionless market does not apply to the housing market. Indeed, most theoretical explanations of the observed relationship between sales and price movements have concentrated on the influence of trading frictions.

Liquidity constraints faced by households may create lock-in costs, reducing the propensity of households to move. During housing market bust, mortgage down payment requirements could make it impossible to sell the old dwelling and buy a new one even if the optimal unconstrained decision was to move. During times of rising housing prices, in turn, credit constraints become less binding augmenting mobility, as argued by Ortalo-Magné and Rady (1999). The impact of liquidity constraints on the co-movement between housing prices and sales is illustrated by the theoretical models of Stein (1995) and Ortalo-Magné and Rady (2006). While Stein highlights the role of downpayment constraints, Ortalo-Magné and Rady (2006) emphasize the interaction between young credit-constrained households with older unconstrained households. In line with these models, Genesove and Mayer (1997) find that seller reservation prices are affected by the loan-to-value ratio; sellers with low equity require longer marketing periods and collect relatively higher prices for their properties than do sellers with more equity. Linneman and Wachter (1989), Zorn (1989) and Engelhardt (1994), in turn, provide empirical evidence for the effect of credit constraints on owner-occupation and mobility rates.

Also the search-theoretic models by Wheaton (1990) and Krainer (2001) suggest that transaction volume (and average selling times) may have a strong contemporaneous relationship with housing price movements even if market participants are fully rational. These models show that financial constraints are not a necessary condition for the liquidity of housing to vary over different states of the world. Wheaton's model, in which capital markets are perfect and all buyers are also sellers, yields a strong theoretical inverse relationship between vacancy rates (and thereby selling times) and prices. Growth in the housing stock (or equivalently increase in the vacancy rate) decreases selling probability, extends average selling times and has a negative impact on prices. The impact of an increase in the number of households is just the opposite. In Wheaton's model also demographic change, in the form of family size, augments transaction volume, which shortens expected selling times and increases housing prices. Also other factors that affect housing demand may influence household mobility and thereby the sales volume and price growth. For instance, the estimations by Berkovec and Goodman (1996) suggest that turnover in the housing market reacts positively to increases in housing demand induced by income growth or lower mortgage rates.

According to Krainer's (2001) model, in "hot" markets prices are rising, average selling times are short and the volume of transactions is higher than the norm. "Cold" markets have the opposite characteristics. As Krainer notes, time-varying liquidity implies that house prices do not vary as much across states of nature as do buyers' valuation of those houses. Krainer also links the correlation between turnover, selling times and price movements to rental market characteristics. The model suggests that if sellers can rent out their unsold houses at fair rates that completely reflect the aggregate state of the economy, then state-varying liquidity and the inverse correlation between expected selling times and price movements disappear. However, the fact that moral hazard and other contracting problems often discourage sellers from renting out their empty houses supports the assertion made by Krainer that hot and cold real estate markets are perfectly consistent with the optimal pricing and buying decisions of forward-looking agents.

Due to the informational inefficiencies in the housing market, there may also be asymmetries in buyers' and sellers' responses because of which the market is likely to exhibit some quantity adjustment. In particular, in search theoretic models where buyers are assumed to respond prior to sellers, sales volume is expected to respond prior to prices. The model by Berkovec and Goodman (1996) assumes that, due to incomplete knowledge of current market conditions, price expectations are backward-looking and prices adapt only gradually toward the equilibrium level. That is, a demand shock has no immediate impact on either buyers' or sellers' price expectations. Instead, initially it only affects the number of buyers on the market and/or their willingness to pay for individual units. This alters the time it takes to sell a house and thereby the number of units sold. Eventually, also selling prices will be affected since sellers' reservation prices are assumed to be negatively related to time on the market. Hence, the model suggests that variations in time to sell transmit changing market conditions to prices. In line with the prediction of the theoretical model, Berkovec and Goodman find empirical evidence for the transaction volume to respond more quickly than prices to shocks. Consequently, they conclude that, for high frequency data, both theoretical and empirical evidence suggest that turnover is superior to price as an indicator of change in housing demand. Also the model by Albrecht et al. (2007) concentrates on the role of sales time in the adjustment process. In the model, buyers and sellers who have not been able to close transactions eventually fall into a desperate state willing to buy at a higher price or sell at a lower price.

Consistent with the assumptions set in the Albrecht et al. and Berkovec-Goodman models, Merlo and Ortalo-Magné (2004) find that a significant fraction of sellers who fail to reach agreement in their first negotiation end up accepting a lower price in the UK. Moreover, according to Merlo and Ortalo-Magné about one-fourth of all sellers make infrequent (typically one) but sizable changes in the list price during the time the house is on the market. These changes are almost always price decreases and are usually substantial. One intuition (suggested by DiPasquale and Wheaton (1996)) behind the lowering of asking prices after being unable to sell is as follows: since most sellers are also buyers, they often temporarily have to own two units. They, therefore, have to carry the housing user cost burden (mortgage payments, maintenance, opportunity cost of equity) of two dwellings. With longer expected selling times, sellers may decide to lower asking prices, preferring to trade sales proceeds for the user cost burden. Home sellers who are not also buyer will be similarly inclined.

Similar to Berkovic and Goodman (1996), Hort (2000) suggests that differences between the adjustment speeds of buyers and sellers to a demand shock might create lead-lag relationship between sales and price movements.[2] According to Hort, the main reason to expect the faster response of buyers is that while buyers' behavior is guided by the direct effect of a demand shock on their individual budget constraint, sellers' decisions are based on estimates of the effect on the distribution of bids. Because information on the aggregate effect is likely to come through more slowly, it is expected that buyers adjust their reservation prices prior to sellers. Furthermore, during the process of search, buyers are more prone than sellers to shop

[2] Also Fisher et al. (2003) explain the positive relationship between price level and market activity within a search model in which changes in seller estimates of property value lag changes in buyer estimates.

around in precisely that segment of the market in which they intend to trade. Therefore, they are likely to be better informed. Moreover, results by Case and Shiller (1988) indicate that in a slow market, sellers' asking prices in particular are rigid and backward looking. Thus, Hort concludes that it seems reasonable to assume that sellers lag buyers in the adjustment of their price expectations.

An additional factor that may induce time variation in sales volume and in average selling times and cause a lead-lag effect between sales and prices is loss aversion. According to the prospect theory (Kahneman and Tversky, 1979), the original purchasing price may be a reference point in the value function of loss-averse agents. The empirical examinations of Genesove and Mayer (2001) and Einiö et al. (2008) support the hypothesis of loss-averse agents in the housing market. Both of these articles find evidence consistent with the idea that sellers are willing to wait longer to get a price at least as large as the original purchasing price. Furthermore, Seiler et al. (2008) find evidence for the sellers to use previous price top as a (false) reference point in the housing market. The existence of loss-averse agents leads to longer average sales times and fewer transactions in down market. In fact, Genesove and Mayer find loss aversion to be a more important factor than liquidity constraints in explaining the price-volume correlation. On the other hand, Case and Shiller (2003) claim that selling times fall in hot markets because overexuberant buyers have irrational house price anticipations, formed by backward-looking expectations. They note that "buyers and sellers in the housing market are overwhelmingly amateurs, who have little experience with trading".

Since losses appear to be calculated in nominal term (Genesove and Mayer, 2001), loss aversion may lead to an even more pronounced drop in transaction volume and increase in sales times in periods with relatively low inflation. During times of high inflation rate, the real housing price level is able to adjust even if nominal prices do not decline. In less inflationary environment, however, real price downward adjustment may well be notably slower causing a greater response in volumes and selling times.

There are a limited number of empirical studies employing dynamic models to investigate whether the predictions of the theoretical models hold true in the actual housing market. Hort (2000) estimates a three-dimensional VAR model including levels of transactions, nominal housing prices and nominal mortgage interest rate using quarterly and monthly panel data from Sweden. Hort assumes that sales and prices are cointegrated and finds empirical support for the assumption. The results provide some support for the hypothesis that sales respond prior to prices to a shock in housing demand (i.e. in mortgage rate). Andrew and Meen (2003a) estimate a two-dimensional "conditional VAR" including housing price growth, turnover rate (transactions / housing stock) and an error-correction mechanism in both equations. The results imply that both prices and transactions respond to housing market disequilibrium, i.e. to the deviation of the desired owner-occupied housing stock from the actual stock. In the model, lagged turnover rate does not enter significantly the equation for price movements. Lagged price change at a four quarter lag, instead, appears to predict transactions with a negative coefficient. In line with the theoretical predictions of the Berkovec-Goodman (1996) model, the empirical examinations by Andrew and Meen indicate that transactions react faster than prices to shocks in the fundamentals and an interest rate shock has a permanent effect on prices but only a transitory impact on turnover.[3]

[3] The use of the turnover *level* in the model yields the zero long-run impact on transactions automatically, though. That is, the model includes implicitly the restriction of zero long-run effect on transactions.

Nevertheless, due to the short-run reaction of the variables, price movements and transactions are positively correlated in the short run.

In summary, the following hypotheses can be derived from the previous literature: 1. There is significant positive (negative) contemporaneous correlation between housing price movements and transaction volume (average selling times) and the correlation is greater at lower data frequencies. 2. Sales volume and selling times respond more rapidly than prices to demand shocks. Therefore, changes in volume and selling times lead price movements. 3. Changes in transaction volume and selling times are temporary whereas changes in prices are permanent (due to the upward sloping supply curve of housing). Furthermore, one additional hypothesis is set in this paper: 4. there is a short-run positive (negative) feedback mechanism from prices to volume (selling times) due to buyers' backward-looking expectations and sellers' slow reaction to market changes. A drop (rise) in the price level affects negatively (positively) the buyers' expectations and reduces (increases) the availability of credit thereby further lowering (increasing) the demand for housing, whereas the sellers do not lower the asking prices at the same pace. That is, it is hypothesized that faster housing price growth predicts greater turnover, i.e. price movements can be used to predict transaction volume and selling times even though prices respond to shocks more slowly.

This paper provides new empirical evidence regarding the co-movement between and dynamics of sales volume and price movements in the housing market. While the previous related literature (Hort, 2000; Andrew and Meen, 2003a) studies the reactions of prices and sales to an interest rate shock only, this article investigates the reactions to income and credit shocks as well. Furthermore, this study appears to be the first one where the dynamics between sales and prices in the housing market are examined rigorously by multiple variable vector-error correction models. In addition, we test if sales volume includes predictive information with respect to housing price movements that is not contained in the fundamental variables driving housing prices, and if price data can be used to estimate better prediction models for transaction volume. The purpose of this article is not to test the alternative theories that try to explain the correlation between price movements and sales. As Meen (2001) notes, any empirical results in this kind of analysis are likely to be consistent with a number of alternative explanations. This is also the case in this study.

3 Long-run model and short-run dynamics

Similar to Andrew and Meen (2003a) the econometric model estimated in this study includes short-run dynamics as well as adjustment towards an estimated long-run equilibrium. A vector error-correction model (VECM), in which the error-correction term is based on a stationary long-run relation between housing price level and its fundamental determinants, is estimated to study the dynamics between sales volume and prices and their responses to demand shocks. In contrast with the Andrew-Meen model, fundamental variables are included in the short-run dynamics of the model as well.

The estimation of the long-run fundamental housing price level is based on the life-cycle model of housing market allowing for the impact of credit constraints (see e.g. Meen, 1990, 2001):

$$P_t = R_t / U_t \quad (1)$$

$$U_t = (1 - T_t)i_t + \gamma - \pi + \delta - \mathrm{E}(P_{t+1} - P_t) + \lambda_t/\mu_c, \tag{2}$$

where P and R stand for the real housing price level and the real imputed rental income, respectively, and U is the housing user cost of capital as a fraction of housing prices. U is determined by the after-tax market interest rate $[(1 - T_t)i_t]$, risk premium (γ)[4] to compensate homeowners for the higher risk of owning than renting, inflation rate (π), property taxes and depreciation of housing (δ), expected real housing appreciation $[\mathrm{E}(P_{t+1} - P_t)]$ and by shadow price of credit rationing constraint (λ_t) divided by the marginal utility of consumption (μ_c). Here, depreciation refers to the maintenance and repair costs that are necessary to maintain constant quality of the structure.

Since the imputed rental income is not directly observable, it is assumed that R is determined by the real per capita income (Y), population (D) and housing stock (H):

$$R_t = \mathrm{f}(Y_t, D_t, H_t). \tag{3}$$

The computation of U includes some notable complications. First, similarly to most other countries, in the Finnish case there is no data concerning λ_t/μ_c. Second, the derivation of expected housing price growth is complicated. Oikarinen (2009b) argues that household debt data are likely to include indirect information regarding the credit constraints, i.e. about λ_t/μ_c. His empirical findings are in line with the argument. Moreover, household debt may well reveal information on income and interest rate expectations as well as on income uncertainty. Therefore, given that there are no direct data on the expected price growth and that the employed expected real price growth series is somewhat arbitrary, loan stock data may bring additional information to the empirical model by telling something about variables that are expected to affect housing demand. Moreover, as illustrated by Stein (1999) and Ortalo-Magné and Rady (1999; 2006), changes in households' access to credit are likely to affect transaction volume in the housing market. This suggests that it is worthwhile to incorporate a credit variable that is likely to strongly correlate with the credit constraints faced by households. Hence, following Oikarinen (2009b), the estimated long-run equilibrium (4) includes a loan stock variable (L) even though debt is not included in the theoretical life-cycle model presented above. Obviously, L presents only a proxy for the correct theoretical concept of λ_t/μ_c.

$$P_t = \beta_1 {*} Y_t + \beta 2 {*} L_t + \beta_3 {*} U_t + e_t. \tag{4}$$

In (4), Y is the aggregate income, i.e. it caters for both population and real per capita income. Betas are the long-run coefficients for the variables explaining long-run development of real housing price level. The error term (e_t) is assumed to be stationary, i.e. the four variables are assumed to form a cointegrating relationship so that housing prices cannot drift away from the price level given by (4) in the long run. Evidently, there are complications in the data as discussed in the data section below. These complications may slightly distort the estimated coefficients.

The estimated long-run relation does not contain any supply side variables. Potential changes in the supply side, such as alterations in the zoning policies, are extremely hard to take into account in an econometric time series analysis. Therefore, it often has to be assumed in

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[4] Meen (1990, 2001) does not include risk premium in the model.

empirical research that there have not been significant changes in the supply side that would affect the long-run relation for housing prices. In this study, a housing stock variable (dwellings per household) was tried in the estimations. The test statistics did not support the inclusion of the stock variable either in the long-run or the short-run model. As the housing stock data embodies notable complications, it is not unexpected that a significant relationship is not found between housing prices and the dwellings per household ratio.[5] Therefore, it is assumed in the econometric analysis that demand (represented by Y, L and U) has driven rental prices and that the housing supply schedule, i.e. the response of housing supply to changes in housing price level, has not notably altered. This assumption is supported by the fact that the estimated long-run model appears to have remained relatively stable through the sample period. Indeed, as the long-run relation includes only demand side variables, recursive test on the stability of the estimated relation may be seen as an indirect test on the stability of the long-run supply curve.

The sales volume is present in the short-run dynamics only. The short-run dynamics and adjustment towards the estimated long-run equilibrium are estimated as a vector error-correction model. In contrast with Andrew and Meen (2003a), fundamental variables are included in the short-run dynamics of the model as well. The fundamentals in the short-run dynamics include Y, L and U. In the Andrew-Meen model, transactions are allowed to be affected only by housing price movements and by market disequilibrium but not directly by market fundamentals. Hort's (2000) VAR model, in turn, incorporates the interest rate as the only fundamental variable in the dynamics.

In this paper, the fundamental variables are allowed to be endogenous. Therefore, it is assumed that the housing market influences the macro variables. This assumption is supported by a great number of empirical studies.[6] However, Y and U are restricted to be weakly exogenous, i.e. they do not react to divergence of the housing price level from its estimated long-run equilibrium. Similarly to Andrew and Meen (2003a), both prices and sales are allowed to respond to deviation from the long-run relation. That is, when prices get too high (low) compared with the fundamentals, sales volume is expected to decrease (increase) and prices are expected to adjust towards the equilibrium. Furthermore, since there are a number of theoretical reasons to expect that credit reacts to the deviation of housing prices from (4) and also empirical evidence supporting the adjustment of credit (Hofmann, 2004; Gerlach and Peng, 2005; Oikarinen, 2009a, 2009b), L too is allowed to respond to disequilibrium.

4 Data

The long-run relation (4) is estimated using quarterly data over 1975Q1-2008Q4. Ideally, the housing price index should be quality adjusted. Unfortunately, hedonic housing price index exists in Finland starting only from 1987. Therefore, similarly to e.g. DiPasquale and Wheaton (1994) and Riddel (2004), an average sales price (per square meter) index and a

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[5] In 1975-1985 the housing stock data are available only at a five-year frequency and in the late sample only at an annual level. Quarterly changes are estimated based on the new housing construction figures.

[6] The results by Oikarinen (2009a, 2009b) provide empirical support for the endogeneity of the fundamentals in the Finnish case.

hedonic price index are joined to have a substantially longer sample period.[7] The use of average transaction prices prior to 1987 may be problematic if the average quality of dwellings sold in different quarters differed notably during the early sample period. It is assumed in the analysis that the price movements displayed by the average sales prices track the true price development well. This assumption is supported by the fact that the difference between hedonic index and average price index is only slight since 1987. The cost of living index is used to deflate the nominal housing price index. The sales volume data (V), in turn, are available since 1987 and are incorporated in the short-run dynamics only. The housing price and sales volume statistics are published by Statistics Finland and are based on transactions of privately financed flats in the secondary market.[8]

The mortgage loan-to-GDP ratio is used as a measure of bank lending (L) in this study. The outstanding mortgage loan stock of households is divided by the nominal GDP to avoid multicollinearity problems in the data.

Income and population are not included in the analysis separately. Instead, it is assumed that the real GDP caters for the influence of both real income per capita and population growth on housing demand. Both loan and GDP data are supplied by Statistics Finland.

As mentioned in section 3, the computation of U is complicated. Because it is extremely difficult to evaluate the rational expectations correctly at a given point of time and because expectations appear to be, to some extent, backward-looking in the housing market, it is understandable that mainly backward-looking expectations have been employed in the empirical literature when estimating the expected housing appreciation (see e.g Poterba, 1992; Englund et al., 1995; Himmelberg et al., 2005; Girouard et al., 2006; Finicelli, 2007). In this study, the expected real appreciation is assumed to be the average real housing price growth in Finland during 1970Q1-2008Q4. The expected inflation, in turn, is estimated as the average quarterly change in the cost of living index during the preceding four quarters. Obviously, the expected housing price growth rate that is employed in the computations is problematic. In particular, the expected real housing appreciation is likely to be time varying in reality. Nevertheless, the computation of expected real appreciation in this article follows the previous literature. Moreover, the purpose of this article is not to compute a perfect measure of U.

The expected real housing price growth rate and expected inflation rate are not the only variables in U that are hard to estimate accurately. In particular, the measurement of the risk premium is complicated. Similarly to the previous literature, γ is assumed to be time-invariant. Following Flamin and Yamashita (2002) γ is set to 2%. Note that through its potential informational content regarding income uncertainty, L may also include information on variation in the risk premium. This further justifies the use of L in the estimated model. The employed interest rate variable is the real after-tax mortgage rate. Because the mortgage rate series, provided by the Bank of Finland, is available only since 1989Q3, the average lending interest rate concerning the whole outstanding loan stock is used to estimate the evolution of the mortgage rate during the early sample.[9] The average lending rate appears to

[7] The results do not notably change even if the average sales price index is employed for the whole sample period.

[8] In 2008, the share of flats of the total Finnish housing stock was 44%.

[9] The source of the lending rate data is Statistics Finland.

proxy well for the mortgage rate.[10] T is the average marginal income tax rate in Finland from 1975 to 1992 and the capital tax rate from 1993 onwards.[11]

Contrary to most of the empirical studies employing U, δ is allowed to vary in time. The proxy for δ is calculated as the average per square meter maintenance costs (including taxes and repairs needed to maintain constant quality) of privately financed flats divided by the average sales price of privately financed flats in the corresponding year. The cost data are at an annual level. Hence, quarterly variation in the maintenance costs is estimated based on the multi-storey housing section of the property maintenance cost indices. Also the cost data are published by Statistics Finland. Finally, λ_t/μ_c is not included in the user cost computations due to the lack of sufficient data. It is argued that L is likely to contain, to some extent, information regarding the credit constraints.

The stock price index (S) is added to the analysis when estimating the prediction models. The findings of Jud and Winkler (2002) and Oikarinen (2009c) indicate predictive power of the stock market performance w.r.t to housing price growth in the US and Finland, respectively. OMX Helsinki CAP index (OMXHCAP) depicts the performance of the publicly traded stocks in Finland.[12]

We employ natural logs of all the variables except for U. Natural logarithm of U is not used, since the estimated user cost is negative during a number of periods. Table 1 presents summary statistics of the differenced series employed in the econometric analysis.[13]

Table 1 Summary statistics of the differenced series for the 1975-2008 period

Variable	**Geometric mean (annualised)**	**Standard deviation (annualised)**	**1st order autocorrelation**
Real housing prices	.013	.062	.628**
Sales volume (1987Q1-)	-.028	.279	-.213*
GDP	.025**	.023	.150
Loan-to-GDP ratio	.040**	.037	.511**
User cost %	.003	.016	.373**
Stocks	.028	.205	.445**

* and ** denote for statistical significance at the 5% and 1% level, respectively.

[10] The correlation coefficient between the average lending rate and the average mortgage rate is .99 from 1989Q3 to 2008Q4. Correlation between the differences is .90.

[11] The source of the national average marginal income tax rate during 1975-1976 is Salo (1990), whereas the data over 1977-1992 is provided by the Finnish Ministry of Finance.

[12] OMXHCAP was formerly called HEX-portfolio index. Prior to 1990 OMXHCAP corresponds to the Unitas index.

[13] Since Wheaton's (1990) search model suggests that demographic change, in the form of family size, augments sales volume, also an average household size variable was tried in the estimations. In contrast with the implications of Wheaton's model, the various information criteria suggest clearly that change in the average household size should be included neither in the VECMs nor the predictions models.

Most of the differenced variables are highly autocorrelated. The negative autocorrelation in the sales series is due to seasonal variation, though. Changes in *V* do not appear to be autocorrelated if the seasonal variation is catered for.

The Augmented Dickey-Fuller (ADF) test implies that real housing prices (*P*), GDP (*Y*), *L* and *S* are non-stationary in levels but stationary in differences (Table 2). This is in line with most of the previous empirical evidence. Based on the ADF statistics, *U* might be stationary. However, since the Johansen test statistics (in section 5) suggest that none of the variables included in the long-run model is stationary, *U* is treated as an I(1) variable.

While Hort (2000) finds sales to be I(1), Andrew and Meen (2003a) handle sales (divided by housing stock) as a stationary variable. In the Finnish case, *V* seems to be stationary. Therefore, *V* is not differenced in the econometric analysis to avoid over differencing and loosing valuable information. Since a constant is needed in the ADF test for *V*, i.e. the ADF test suggests that *V* is trend stationary, a detrended version of *V* is used in the forthcoming empirical analysis. The growing trend in sales most likely reflects the fact that the overall stock of flats has increased thereby leading to a greater number of sales even if the turnover rate has remained constant. Naturally, it would be optimal to employ the turnover rate (sales / stock) in the analysis. Unfortunately, there does not appear to be a long enough time series regarding the number of privately financed flats in Finland. Anyhow, the fact that *V* seems to be trend stationary implies that a deterministic trend can cater for the growing trend in sales caused by the growth of the stock. This is unsurprising, since the overall Finnish housing stock has grown quite steadily.[14]

Table 2 Augmented Dickey-Fuller test results

Variable	Level (lags)	Difference (lags)
Housing price	-1.85 (5)^{c,s}	-3.73** (4)
Sales volume (1987Q1-)	-3.34* (4)^{c,s}	-3.31** (3)^{s}
GDP	-1.41 (3)^{c}	-2.47* (2)
Loan-to-GDP ratio	.18 (3)^{c,s}	-3.83** (2)^{s}
User cost %	-1.94 (3)	-4.47** (2)
Stocks	-1.35 (5)^{c}	-4.88** (4)

* and ** denote for statistical significance at the 5% and 1% level, respectively. Critical values at the 5% and 1% significance levels are -1.95 and -2.60 if constant is not included and -2.89 and -3.51 in the case where constant is present. The number of lags included in the ADF tests is decided based on the general-to-specific method. A constant term (^{c}) is included in the tested model if the series clearly seem to be trending or if the ADF test without the constant term suggests that the series are exploding. In addition, three seasonal dummies (^{s}) are added to the test if recommended by the F-test.

The theory expects sales volume and price movements to be positively correlated. Table 3 reports a correlation matrix regarding contemporaneous quarterly correlations between the variables. Because data on transactions exist only since 1987, all the correlation figures are based on a sample period over 1987-2008. Expectedly, the contemporaneous correlation between sales and housing price movements is large (.7) and statistically highly significant. There is no correlation between ΔV and ΔP, however. In contrast with Follain and Velz

[14] A constant and a deterministic trend together explain over 99% of the annual variation in the Finnish housing stock during 1987-2008.

(1995) and Hort (2000), in the Finnish case there is no evidence of negative correlation between sales and price level at the quarterly frequency.

Table 3 Contemporaneous quarterly correlations between differenced series over 1987-2008

	ΔP	ΔV	ΔY	ΔL	ΔU	ΔS	V	P
ΔP	1							
ΔV	.01	1						
ΔY	.70***	.05	1					
ΔL	.05	.03	-.27**	1				
ΔU	-.49***	.01	-.46***	.04	1			
ΔS	.46***	.33***	.36***	-.20*	-.24**	1		
V	.70***		.51***	.30***	-.41***	-.53***	1	1
P		-.17*					.10	1

*, ** and *** denote for statistical significance at the 10%, 5% and 1% level, respectively.

Table 4 Autocorrelations and cross-autocorrelations between housing price change and sales volume over 1987-2008

	Lagged housing price change							
	-1	-2	-3	-4	-5	-6	-7	-8
Price change	.73***	.60***	.42***	.40***	.21*	.08	.02	-.04
Sales volume	.53***	.39***	.25**	.16	.05	.02	-.16*	-.09
	Lagged sales volume							
	-1	-2	-3	-4	-5	-6	-7	-8
Price change	.69***	.56***	.50***	.44***	.26***	.14	.06	-.01
Sales volume	.70***	.53***	.49***	.37***	.10	.01	.-11	-.22**

*, ** and *** denote for statistical significance at the 10%, 5% and 1% level, respectively.

Table 4 shows that, in addition to the positive short-term autocorrelations, the cross-autocorrelations between price movements and sales are large and significant up to several quarter lags. In particular, the correlations between housing price movements and lagged sales volume are great and significant up to five quarters. Moreover, Figure 1 shows that peaks (bottoms) in the transaction volume have preceded peaks (bottoms) in the price level. To see this phenomenon more prominently, a Hodrick-Prescott filtered sales volume is shown in the middle part of the figure. The indices in Figure 1 are not in the natural log form.

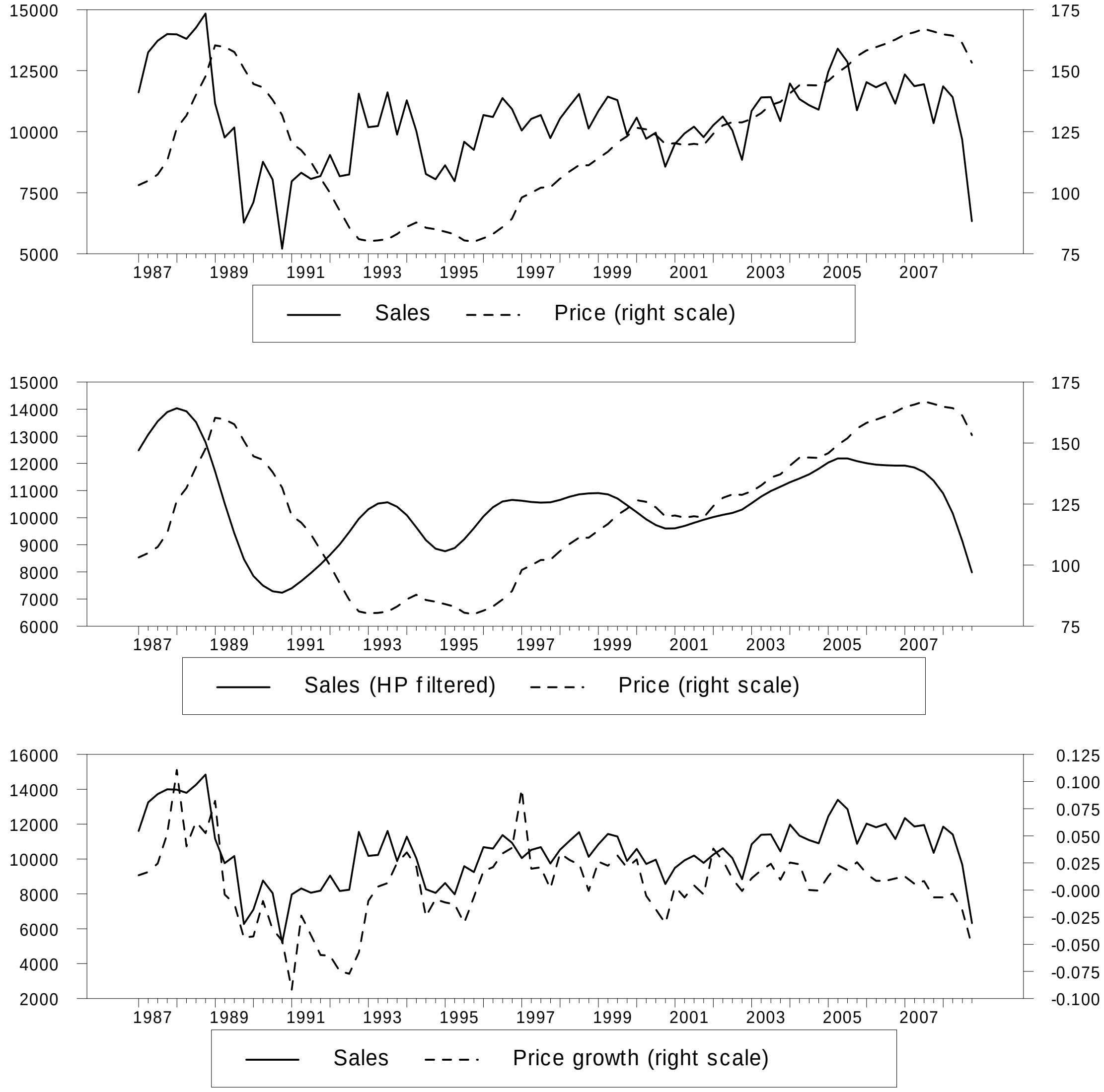

Figure 1 Real housing price index, sales volume and housing price growth over 1987-2008

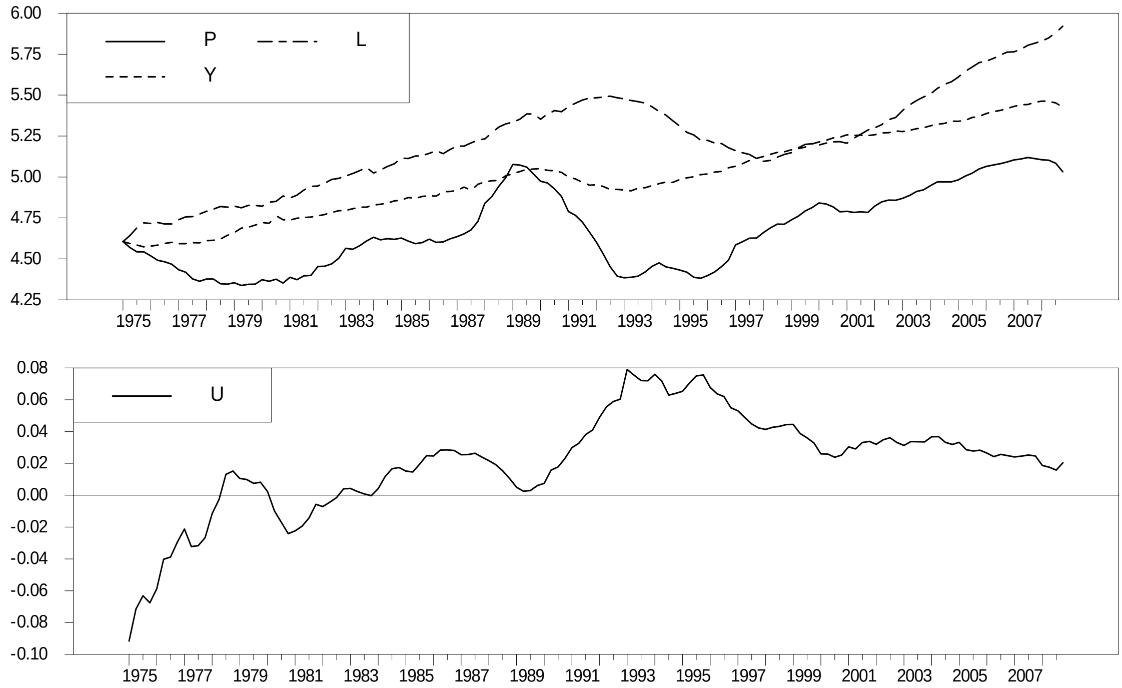

Figure 2 Real housing price index (P), real GDP index (Y), loan-to-GDP ratio index (L) and user cost of housing (U) over 1975-2008

The evolution of the fundamental variables and of housing prices over the sample period is shown in Figure 2. Due to the low, administratively controlled, lending rates together with foreign capital controls, credit rationing was effective in Finland up to the mid-eighties. Because of the controls, the real after-tax mortgage rate and the user cost of capital were negative for sustainable periods in the 1970s and early 80s. A financial market deregulation took place in Finland in the late 1980s. The financial deregulation resulted in looser liquidity constraints, induced a huge growth of credit and led to a housing market boom and finally to a housing price overshot. This can be well seen in Figure 2. While the sales volume peaked in 1988Q1, prices continued to rise until 1989Q1 (Figure 1). Eventually, prices started to decline sharply. The drastic drop in housing prices was emphasized by the rising real interest rates and the deep recession of the Finnish economy.[15]

While the rapid housing price drop did not stop until 1993, sales started to notably increase already in 1990. Since 1996 the real price level has substantially increased (92% between 1995Q4-2008Q4). Also V trended up until 2005. In 2007, however, sales started to decline and in 2008 the drop in V was greatly emphasized by the global financial crisis. Again, the drop in V preceded a decrease in P. Similar to the late 1980s, the lag appears to have been approximately one year.

Note that according to Oikarinen (2009a, 2009b) the financial market liberalization induced a structural break in the short-run dynamics between P and L. Hence, it would be reasonable to estimate the short-run dynamics using data from the late 1980s onwards even if the sales data covered a longer sample period.

[15] For more details about the boom-bust cycle see e.g. Honkapohja and Koskela (1999).

5 Econometric methodology

In the econometric section, the Johansen test for cointegration is employed to investigate the stationarity of the long-run relation (4) and to estimate the long-run parameters. The cointegration analysis works as a specification test for (4). If cointegration between the variables in the long-run model could not be detected, then e_t would not be stationary and the model would be misspecified. Since *Y* and *U* are assumed to be weakly exogenous, the methodology and asymptotic tables presented in Harbo et al. (1998) are applied to increase the power of the Johansen Trace test. Harbo et al. suggest that in the case of weakly exogenous variables the Trace test should be based on the following VECM:

$$\Delta X_t = \alpha(\beta_X', \beta_1)(X'_{t-1}, t)' + \alpha\beta_Z' Z_{t-1} + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1}\Delta X_{t-k+1} + \gamma_1 \Delta Z_{t-1} + \ldots + \gamma_{k-1}\Delta Z_{t-k+1} + \mu + \Psi D_t + \varepsilon_t, \quad (5)$$

where X_t is a 2-dimensional vector of the adjusting stochastic variables (*P* and *L*), ΔX_t is $X_t - X_{t-1}$, Z_t is a vector of the weakly exogenous stochastic variables (*Y* and *U*), ΔZ_t is $Z_t - Z_{t-1}$ and $t = 1,\ldots,T$. Γ_i and γ_i are 2 x 2 matrices of coefficients for the lagged differences of the stochastic variables at lag i, k is the maximum lag, i.e. the number of lags included in the corresponding vector autoregressive model in levels, μ is a 2-dimensional vector of intercepts, D_t is a (s-1)-dimensional vector of centered seasonal dummies (in this study s = 4, since quarterly data is used), Ψ is a 2 x (s-1) coefficient matrix and ε_t is a 2-dimensional vector of of white noise error terms. Furthermore, α is a 2-dimensional vector of the speed of adjustment parameters, while $(\beta_X', \beta_1)(X'_{t-1}, t)' + \alpha\beta_Z' Z_{t-1}$ forms the cointegrating relationship and β_X' and β_Z' include the parameter estimates for the long-run relation (4). The cointegrating relation in the tested model also incorporates a deterministic trend (t).

The maximum lag is selected based on the Hannan-Quinn information criteria (HQ) together with the LR(1) and LR(4) tests for residual autocorrelation. Furthermore, since some of the series seem to exhibit seasonal variation, the need for seasonal dummies is detected based on HQ. The asymptotic tables by Harbo et al. are compared with the estimated Trace test statistics to check the existence of a cointegrating vector between the variables. The exclusion of the stochastic variables from the long-run relation is tested by the Bartlett small-sample corrected Likelihood Ratio (LR) test presented in Johansen (2000). Finally, the stability of the estimated long-run relation is examined employing a recursive estimation analysis explained in Juselius (2006).

Based on the estimated long-run relation, VECMs are estimated to study the dynamics more carefully. In addition to the four variables in the long-run model, sales are added in the VECMs. Sales data appear in the models in levels, since sales volume seems to be stationary. The VECMs are used to study Granger causalities between the variables and to estimate the impulse responses of prices and sales to demand shocks. Granger causalities are investigated by a standard F-test and the Choleski decomposition is utilized to derive the impulse responses. While the cointegration test is conducted using data for the 1975Q1-2008Q4 period, the short-run dynamics in the VECMs are based on data from 1987Q1 onwards.

In addition, we estimate short-run prediction models and compute out-of-sample forecasts to further evaluate the predictive properties of housing prices and sales with respect to each other. The prediction models are compared based on the root mean squared error (RMSE) statistics. The number of lags included in the VECMs and the variables included in the prediction models are decided based on Schwarz Bayesian Information Criteria (SBC).

6 Empirical results

In this section, an estimate for the long-term "equilibrium" housing price level is derived first. Then a model including the short-run dynamics between prices, sales and the fundamental variables and the reaction of prices and sales to "disequilibrium" is estimated. While the long-run relation is based on data over 1975Q1-2008Q4, a sample period from 1987Q1 to 2008Q4 is used in the estimation of the short-run dynamics.

As housing prices, GDP and the loan-to-GDP ratio seem to be I(1), cointegration analysis is used to test if there is a stationary long-term relationship between housing prices and the fundamentals. The Johansen Trace test statistics (reported in Table 5), based on a vector error-correction model in which *Y* and *U* are restricted to be weakly exogenous, suggest that there are two stationary linear vectors in a system including *P*, *Y*, *L* and *U*. The LR test statistics suggest that the deterministic trend in the long-run model can be excluded. While one of the stationary vectors can be regarded as the long-run relation between housing prices and the fundamentals, the other stationary vector does not appear to include *P*.[16] Therefore, for now on only the vector including housing prices is analyzed and the LR test results reported in Table 5 are based solely on this relation. Note that *U* does not enter the estimated long-run model for housing prices.

Table 5 Cointegration test statictics

H_0 (rank)	Trace test value (p-value)
R = 0	46.0
R ≤ 1	20.7
P-value in the test for exclusion of *t*	.46
P-value in the test for exclusion of *t* and *U*	.39
Estimated long-run relation (standard errors in parenthesis):	
$P = .094 + \underset{(.168)}{.482} {*} Y + \underset{(.138)}{.200} {*} L$	

The tested model includes four stochastic variables (*P, Y*, *L*, *U*) and incorporates four lags in differences, three centered seasonal dummies and a deterministic trend in the long-run dynamics. 95% quantiles are 35.5 (R = 0) and 17.9 (R = 1). Exclusion of the variables is tested by the Bartlett small sample corrected LR test by Johansen (2000).

The estimated relation suggests that one percent increase in GDP leads to approximately .5% higher housing prices. The coefficient of the mortgage-to-GDP ratio is somewhat smaller (.2). The insignificance of *U* in the long-run model is not totally unexpected. It is reasonable to assume that *U* is mean-reverting. As Goodhart and Hofmann (2007) state, real interest rate is usually considered to be mean-reverting, and it is the inflation and interest rate movements that are major factors driving the user cost variable employed in this study. Mean-reversion in *U* is supported also by the ADF test statistics.[17] If *U* is indeed mean-reverting, then the

[16] P-value in the joint test for exclusion of trend and *U* from the first cointegrating vector and for exclusion of *P* and *U* from the second cointegrating vector is .41. This result also implies that *U* alone does not form a stationary vector. Thus, the LR test statistics are in contrast with the ADF test results regarding the stationarity of *U*.

[17] Also Meen (1996) argues that the user cost is stationary.

housing demand of forward-looking agents with long planned holding periods of housing should not react strongly to changes in it. In the presence of large transaction costs and relatively low liquidity this is likely to be true at least concerning the owner-occupants. This may explain the insignificance of *U* in the estimated long-run relation. In line with this argument, Shiller (2007) writes: "*People's opinions about long-term decisions, notably how much housing to buy and what is a reasonable price to pay, change in the short term only because their opinions about long-term change*". Note also that the findings by Also Aoki et al. (2002), Gerlach and Peng (2005), Elbourne (2008) and Oikarinen (2009a), according to which the effect of interest rate changes on housing prices is insignificant or relatively small, are in line with the insignificance of *U* in the long-run model estimated in this paper.

Figure 3 shows that *P* has substantially deviated from the estimated long-run relation during the sample period. Nevertheless, the relation has been relatively stable according to the recursive test (see Figure A1 in the Appendix). The price level was relatively close to the long-run relation until late 1987. The financial market liberalization resulted in a boom in the housing market and in 1989Q1 the real housing price level peaked being over 40% above the long-run relation. Eventually, the prices started to decline drastically and housing prices overreacted downwards in the early and mid 1990s. Three years after the price peak *P* was almost 30% below the estimated long-run level. In 1996 the real housing price level started to rise again. Since then *P* has increased by 90% (the situation in 2008Q4). In 2008Q4 the price level was only 1% over the long-run relation after being some 10% above the relation during the preceding couple of years.

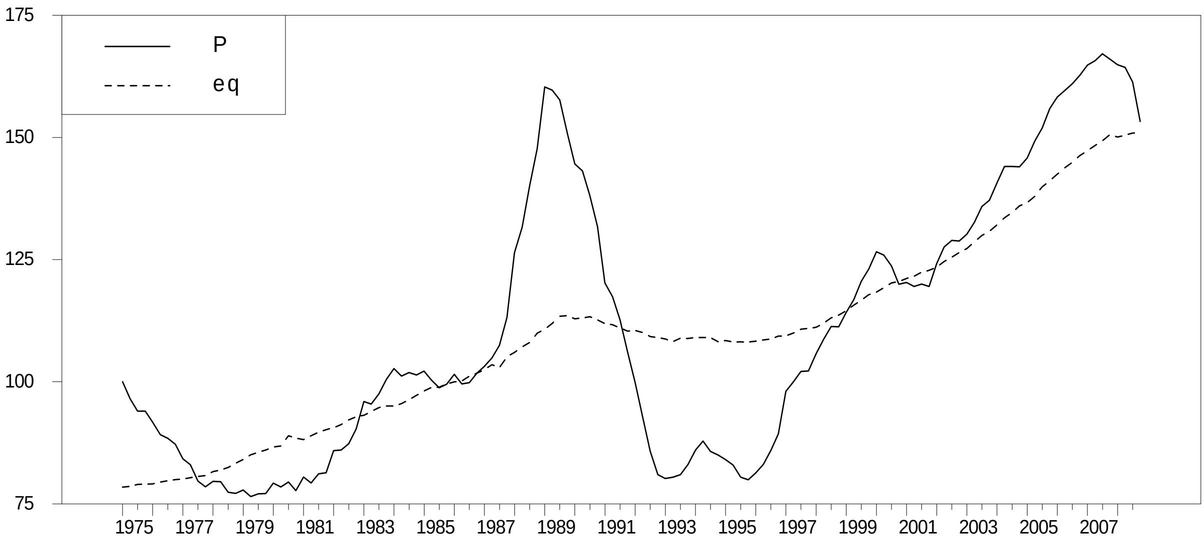

Figure 3 Housing price index (H) and fit from the estimated long-run relation (EQ)

Granger causalities between transaction volume and price changes are investigated based on two VECMs, a pairwise model and a model that includes also the fundamental variables in the short-run dynamics. Deviation from the long-run relation reported above is used in the error-correction mechanism. The Granger non-causality test results are presented in Table 6. Both price growth and sales are highly predictable and Granger cause themselves. In line with the theoretical considerations by Berkovec and Goodman (1996) and Hort (2000), V clearly Granger causes ΔP. Based on the pairwise model there is no feedback from price movements to sales. However, the economically large coefficient of .75 on ΔP in the sales equation in the multiple variable model suggests that there may actually be notable feedback from price movements to V even though the p-value is as large as .20.

Furthermore, the relatively large speed of adjustment parameters (-11% and -10% in the pairwise and multiple variable models, respectively) estimated for V imply that sales too react to deviation of housing prices from the long-run relation. That is, congruent with the empirical findings by Andrew and Meen (2003a) prices appear to Granger cause V through the error-correction mechanism. Also in line with Andrew and Meen, the speed of adjustment parameters suggest that sales adjust more rapidly than prices to shocks. Note, however, that although being economically large, the speed of adjustment parameters of V are statistically insignificant due to their large standard deviations.[18] It also has to be noted that the p-values in the equations for volume should be taken only as suggestive, since the residuals appear to exhibit non-normality.

Table 6 P-values in the Granger non-causality tests

Dependent Variable	Explanatory variable						
	ΔH	V	ΔY	ΔL	ΔU	eqe	Adj. R^2
Pairwise model							
Δ Price	**.00**	**.00**				**.01**	.77
Volume	.92	.00				.17	.56
Multiple variable model							
Δ Price	**.00**	**.00**	.44	.64	.14	**.09**	.73
Volume	.20	**.00**	**.06**	.75	.95	.39	.59

The reported p-values are based on a standard F-test. Statistically significant values at the 10% level are bolded and eqe denotes equilibrium error, i.e. deviation from the long-run relation. The pairwise model includes two lags in differences while the six variable model includes one lag in differences. Both models include three seasonal dummies. Due to heteroscedasticity of the residuals, the p-values in the ΔP equation in the multiple variable model are based on a covariance matrix that is computed allowing for heteroscedasticity as in White (1980).

The multiple variable VECM is used to study the response speeds of volume and prices to changes in the fundamentals more rigorously (the model is summarized in Table A1 in the Appendix). Given the economically significant size of the speed of adjustment parameter of V and the results of Andrew and Meen (2003a) that support the reaction of sales to housing price disequilibrium, V is allowed to react to the deviation from the estimated long-run relation. Thus, it is assumed that when prices get too high (low) compared with the

[18] The more parsimonious prediction models present stronger support for the reaction of transaction volume to disequilibrium.

fundamentals, sales volume decreases (increases). Also *P* and *L* adjust towards the long-run relation in the model, the speed of adjustment parameters being -4.1% for prices and 3.7% (and highly significant) for *L*.

Figure 4 shows the impulse responses up to 40 quarters from the initial shock. The impulse responses are based on Choleski decomposition using the following ordering: *Y*-*L*-*U*-*P*-*V*. It is therefore assumed that GDP does not contemporaneously respond to innovations in any of the other variables, but may affect all the other variables within the quarter. The ordering also reflects the common assumption that the interest rate changes are transmitted to the economy with lag (through *U* in the estimated model). However, shocks to all the fundamental variables, including *U*, are allowed to contemporaneously affect prices and sales. Since the theory and the empirical results suggest that changes in the transaction volume lead (i.e. have a delayed impact on) prices, shocks to *P* are allowed to have an immediate impact on *V* but not the other way round. Anyhow, the impulse responses are robust to the ordering between *P* and *V*.

The response of prices, and also of sales to somewhat lesser extent, to demand shocks appears to be cyclical. In line with the theoretical consideration of Berkovec and Goodman (1996) and Hort (2000), *V* reacts more rabidly than *P* to the shocks. Within a quarter from a credit shock, sales substantially increase while the price response is only small. As the price level gradually increases and overshoots, sales volume declines. After around two years, also prices start to decrease. Eventually, *P* converges towards a level that is .7% greater than its initial level. A similar pattern but with negative signs can be seen in the case of a user cost shock. Again, there is an initial strong response of *V*, whereas *P* reacts sluggishly and overshoots before converging towards its new long-run level. In both cases the sales volume seems to be smaller than the initial one for some while after the overshot in the price level. This is mainly due to the reaction of *V* to deviation of the price level from its long-run relation, i.e. the level of sales is relatively low when prices are “too” high. Together, the immediate strong response of sales volume and the slow reaction of prices suggest that the sluggish price adjustment is due to the gradual change in the sellers’ reservation prices after a shock.

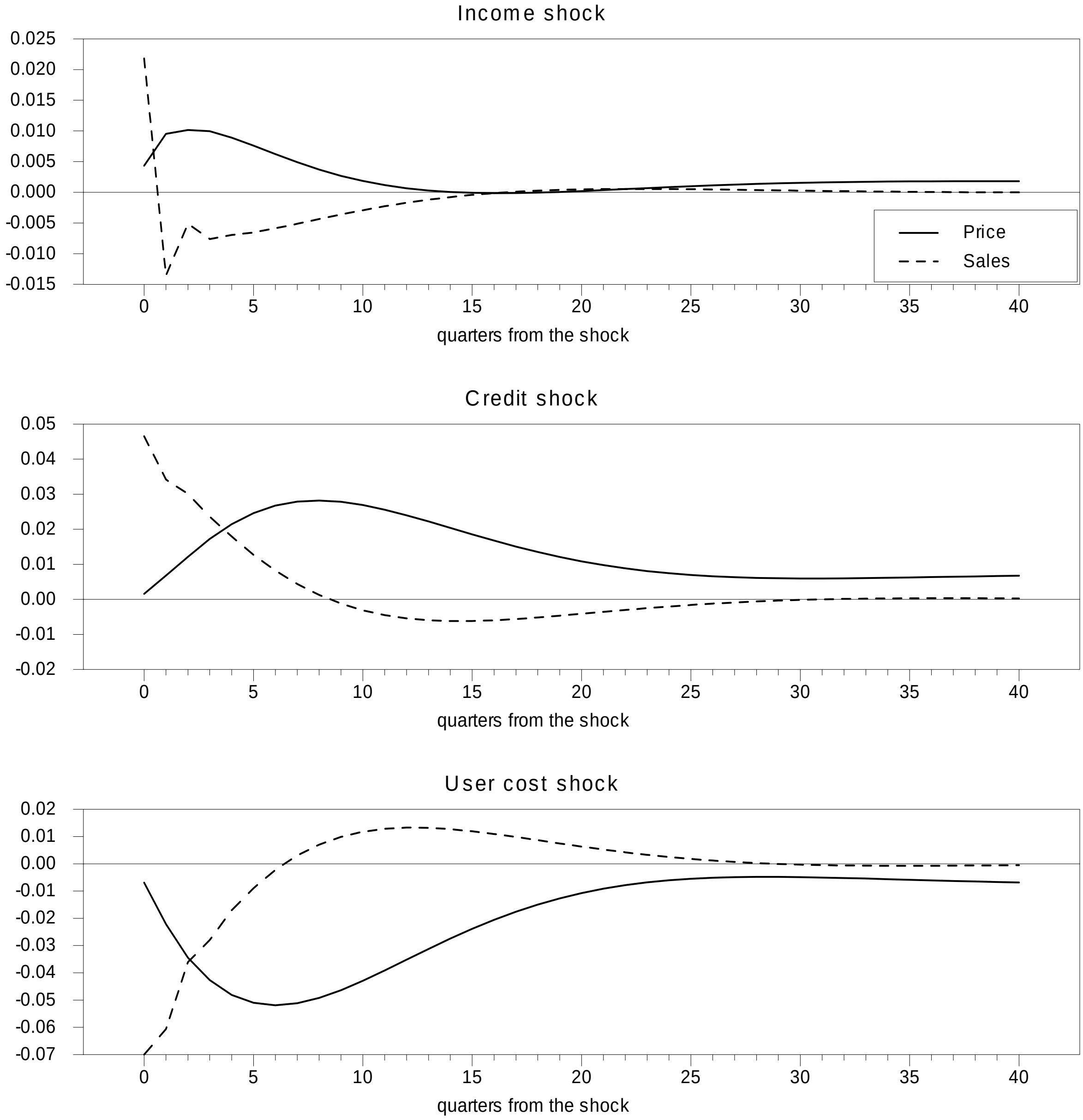

Figure 4 Impulse responses of housing prices and sales volume to one unit shocks in the fundamental variables

The response of V to an income shock is not as straightforward to interpret. The estimated initial reaction of V to a positive GDP shock is large and positive, after which sales notably decrease. The decline is due to three factors: the immediate increase of the price level above its new long-run level, the negative short-run impact of a shock to Y on L, and the negative coefficient on the lagged GDP growth in the equation for V.

According to the estimated impulse responses, the adjustment patterns of sales and prices after demand shocks create strong co-movement between sales and price growth. After a credit shock and a user cost shock, the contemporaneous correlation between sales and price movements is .85 and .90, respectively. While both of these correlations are highly significant, the correlations between price level and sales (-.12 and -.01) are not statistically significantly different from zero. In the case of an income shock, however, the negative correlation between the responses of the levels (-.48) as well as the positive correlation between price movements and sales (.31) are statistically significant.

The relatively small long-term impacts of income and user cost shocks on housing prices may be due to the inclusion of *L* in the model. As explained in section 3, *L* is likely to incorporate information about interest rate and income expectations thereby diminishing the influence of the current level of interest rates (which materializes through the user cost variable) and of *Y*.

The reaction of *V* to an interest rate shock (i.e. user cost shock) is somewhat different from that reported by Hort (2000) and by Andrew and Meen (2003a).[19] In particular, sales react notably faster than based on the Hort and Andrew-Meen models. In fact, the maximum effect of a user cost shock on sales takes place immediately after the shock. Given that Andrew and Meen do not include fundamental variables in the short-run dynamics of their model, the difference of the reaction speeds estimated in this study compared with those reported by Andrew and Meen is not unexpected.

Note that the long-run response of *V* to the shocks is zero by construction, since sales volume is assumed to be stationary.[20] This is in line with the argument by Berkovec and Goodman (1996) according to which shocks induce only temporary changes in transactions. The stationarity of *V* is supported by the ADF unit root test and by the stationarity of the residuals derived from the equations estimated for *V*.

To study the predictive power of prices and sales w.r.t each other, we estimate more parsimonious prediction models and compare their out-of-sample forecasts. Three different models for *V* are estimated: one with lagged price change and disequilibrium, one without lagged ΔP but including lagged disequilibrium and one excluding all the variables that incorporate housing price data. Similarly, three models are estimated for ΔP, but now with lagged *V* only in one model. The inclusion of the other variables in the models is based on SBC. The goodness of the models is checked by the root mean square errors (RMSE) of the out-of-sample forecasts.[21] It is only the relatively short-run predictive abilities of the models that are tested. The accuracy of one quarter ahead out-of-sample forecasts are evaluated employing models that are based on quarterly data, whereas annual data is used to estimate the models based on which one year ahead forecasts are derived. Only one period ahead forecasts are studied here to avoid the complications of having to predict the evolution of the fundamental variables. The set of prediction models together with their RMSEs are presented in Tables A2 through A5 in the Appendix. Note that it is the non-detrended sales that is used in the predictions models and a trend is included in the models if suggested by SBC. The annual models utilize the long-run coefficients reported in Table 5.

The main implication of the short-term prediction models is that the predictive power of the lagged observations of price changes and sales overwhelms any direct predictive ability of the lagged changes in the fundamentals. This is clearly the case according to SBC. The best

[19] Neither Andrew and Meen nor Hort examine the reactions of prices and sales to shocks in the other fundamental variables.

[20] The same applies also to the Andrew and Meen (2003a) model.

[21] The quarterly out-of-sample forecasts are derived for the 2006Q1-2008Q4 period while the annual out-of-sample forecasts are computed for the 2004-2008 period.

annual out-of-sample forecasting ability regarding housing price growth is given by a model that includes lagged credit change instead of lagged sales, however. At the quarterly level, instead, lagged data on sales improves the forecasting performance of a model for ΔP. Moreover, the inclusion of sales data removes the heteroscedasticity and non-normality problems in the quarterly model for housing price growth. It appears that, if sales data is not included in the model, also the stock market performance as well as changes in Y and in L exhibit predictive power w.r.t. quarterly housing price growth. Regarding the models for transaction volume, lagged price change has slightly improved the quarterly out-of-sample forecasts. At the annual level, instead, the most accurate out-of-sample predictions have been given by the model with only lagged sales and quarterly dummies.

Appendix 2 reports similar annual prediction models regarding the Helsinki metropolitan area housing market. The implications of these models are in line with the nationwide models. In fact, in the HMA case the results are even more confirming concerning the predictive power of sales volume w.r.t. housing price growth.

7 Summary and conclusions

This article provides new empirical evidence on the co-movement between and dynamics of prices and sales volume in the housing market using quarterly data over 1975-2008 from Finland. A particular emphasis is given to examining the reaction speeds of prices and sales to demand shocks. While the previous related literature studies the reactions of prices and sales to an interest rate shock only (see Hort, 2000; Andrew and Meen, 2003a), this article investigates the reactions to income and credit shocks as well. Furthermore, this study appears to be the first one where the dynamics between sales and prices in the housing market are examined rigorously by a multiple variable vector-error correction model.

The purpose of the study is not to test the alternative theories that try to explain the correlation between price movements and sales. Nevertheless, in line with the theoretical considerations of Berkovec and Goodman (1996) and Hort (2000), sales volume is found to Granger cause price movements and the response of prices to demand shocks is found to be substantially slower than that of sales. That is, in line with the suggestion by Berkovec and Goodman, the findings indicate that turnover is superior to price as an indicator of change in housing demand. The Berkovec-Goodman model as well as Hort's considerations suggest that the leading role of sales is due to the slower response of sellers than buyers to shocks in the market fundamentals. The findings in this paper are consistent with those considerations; the immediate strong response of sales volume suggests that the sluggish price adjustment is due to the slow reaction of the sellers' reservation prices to demand shocks. Interestingly, the estimated reaction speed of sales to shocks in the fundamentals is somewhat different from those presented in the previous empirical literature. In particular, the maximum effect of GDP, credit and user cost shocks on sales takes place within a quarter from the shock. Therefore, sales react faster than based on the empirical models by Hort (2000) and Andrew and Meen (2003a).

While price movements and sales Granger cause themselves and sales also Granger cause price growth, only weak support for positive feedback from price movements to sales is found. Importantly, the predictive power of the lagged observations of price changes and sales overwhelms any direct predictive ability of the lagged changes in the fundamentals that drive housing demand.

Although the co-movement and lead-lag relation between sales and prices in the housing market are likely to be, to a large extent, an outcome of market frictions, the volatility of transaction volume and average selling times may have some beneficial consequences on the overall economy. In particular, the adjustment of volume and selling times may decrease the observed volatility of housing prices. According to Krainer (2001), for instance, state-varying liquidity implies that house prices do not vary as much across states of nature as do buyer valuations of those houses. Genesove and Mayer (2001), Fisher et al. (2003), and Goetzmann and Peng (2006) report some findings in line with such smoothing impact. Given the important role that housing wealth seems to play on aggregate demand through its impact on construction activity, household consumption and on the credit market, smaller perceived volatility might somewhat diminish cycles of the overall economy. This is the case even if potential buyers' valuation of housing, on average, differs from the perceived prices, since it is generally the perceived wealth that affects household consumption.

In Finland, mortgages are typically not tied to a specific dwelling. Instead, one can sell a dwelling and buy a new one without having to apply for a new mortgage. Therefore, the effect of liquidity constraints on the transaction volume is not likely to be as strong in Finland as in the countries where mortgages are usually fixed to a particular dwelling. Nevertheless, the sales volume has been highly volatile and the co-movement between sales and price movements has been strong in Finland. If tighter liquidity constraints obtained during housing market busts, the volatility and co-movement would most likely be even greater.

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Appendix 1

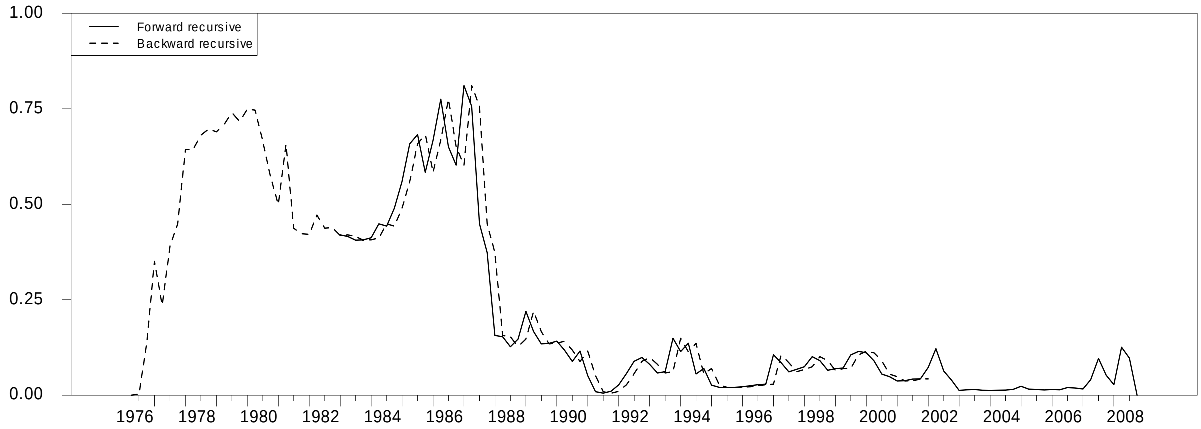

Figure A1 Plot of the recursive and backwards recursive Max Test statistics (in the R-form) of constancy of the estimated long-run relation scaled by the 5% critical value

Table A1 Estimates of the multiple variable VECM for real housing prices movements and for sales volume

	ΔP		V	
Constant	.018***	(.006)	.097***	(.033)
ΔP_{t-1}	.373***	(.121)	.784	(.604)
V_{t-1}	.096***	(.020)	.805***	(.123)
GDP_{t-1}	.243	(.315)	-3.34	(1.76)
ΔL_{t-1}	.113	(.239)	-.370	(1.11)
ΔU_{t-1}	-.641	(.429)	.222	(3.79)
$Q2$	-.022***	(.006)	-.079**	(.039)
$Q3$	-.025***	(.005)	-.071*	(.037)
$Q4$	-.028***	(.006)	-.182***	(.038)
$p_{t-1}-p^e_{t-1}$	-.041	(.025)	.094	(.114)
Adjusted R^2	.73		.58	
J-B	.06		.00	
LM(1)	.46		.68	
LM(2)	.13		.05	
LM(4)-heter	.00		.29	

Standard errors in parenthesis. *, ** and *** denote for statistical significance at the 10%, 5% and 1% level, respectively. *Q2*, *Q3* and *Q4* are three centered seasonal dummies. The standard errors in the equation for ΔP are based on a covariance matrix that is computed allowing for heteroscedasticity as in White (1980). J-B stands for the Jarque-Bera test for residual normality, LM(1) and LM(2) are the Lagrange-multiplier tests for residual autocorrelation at lag lengths one and two, and LM(4)-heter is the fourth order Lagrange-multiplier test for heteroscedasticity in the residuals. Due to the non-normality of the residual in the equation for *V*, the statistical significances in the equation should be taken cautiously.

Table A2 Prediction models for quarterly sales volume

	ECM1		ECM2		ECM3	
Constant	1.14	(.454)	.958	(.328)	1.02	(.326)
Trend						
Δp_{t-1}	.311	(.535)				
V_{t-1}	.760	(.102)	.801	(.074)	.787	(.074)
Q2	-.061	(.037)	-.059	(.037)	-.059	(.037)
Q3	-.059	(.037)	-.060	(.037)	-.060	(.037)
Q4	-.166	(.037)	-.169	(.037)	-.169	(.037)
$p_{t-1}-p^{e}_{t-1}$	-.109	(.078)	-.105	(.078)		
Adj. R^2	.59		.59		.59	
SBC	42.7		38.6		36.1	
			p-values		p-values	
J-B	.00		.00		.00	
LM(1)	.58		.41		.66	
LM(2)	.07		.03		.14	
LM(4)-heter	.19		.12		.13	
RMSE	.130		.131		.132	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. Statistical significances are not denoted, since the residual normality assumption is not fulfilled in the models.

Table A3 Prediction models for annual sales volume

	ECM1		ECM2		ECM3	
Constant	2.85	(2.38)	4.61**	(1.85)	10.3***	(1.44)
Trend					.013***	(.003)
Δp_{t-1}	-.378	(.325)				
V_{t-1}	.732***	(.224)	.565***	(.174)	.419**	(.153)
V_{t-2}					-.417***	(.134)
$p_{t-1}-p^{e}_{t-1}$	-.202	(.170)	-.277*	(.158)		
Adj. R^2	.59		.34		.66	
SBC	-20.6		-22.6		-35.9	
			p-values		p-values	
J-B	.70		.70		.98	
LM(1)	.04		.09		.08	
LM(2)	.06		.14		.21	
LM(4)-heter	.53		.57		.54	
RMSE	.105		.109		.103	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. ***, **, * denote for statistical significance at the 10%, 5% and 1% level, respectively.

Table A4 Prediction models for quarterly price growth

	ECM1		ECM2		ECM3	
Constant	-.233***	(.084)	.008	(.006)	.016***	(.005)
Δp_{t-1}	.424***	(.104)	.444***	(.127)	.669***	(.076)
Δp_{t-2}	.260**	(.088)	.206**	(.094)		
V_{t-1}	.127***	(.017)				(.074)
V_{t-2}	-.069***	(.021)				
ΔGDP_{t-1}			.597	(.444)		
ΔL_{t-1}			.511**	(.247)		
ΔS_{t-1}			.058**	(.022)	.074***	(.024)
$p_{t-1}-p^{e}_{t-1}$	-.033***	(.011)	-.063**	(.028)		
Q1	-.033	(.006)	-.018**	(.007)	-.024***	(.007)
Q2	-.032	(.005)	-.021***	(.006)	-.021***	(.007)
Q3	-.032	(.005)	-.020***	(.006)	-.019***	(.006)
Adj. R^2	.77		.67		.61	
SBC	-289.3		-257.3		-255.9	
			p-values		p-values	
J-B	.32		.00		.12	
LM(1)	.50		.04		.09	
LM(2)	.67		.02		.08	
LM(4)-heter	.16		.00		.13	
RMSE	.013		.016		.015	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. ***, **, * denote for statistical significance at the 10%, 5% and 1% level, respectively. The standard errors in ECM2 are based on a covariance matrix that is computed allowing for heteroscedasticity as in White (1980). Due to the non-normality of the residual, the statistical significances in ECM2 should be taken cautiously.

Table A5 Prediction models for annual price growth

	ECM1		ECM2		ECM3	
Constant	-5.16***	(1.11)	-.025	(.021)	-.058**	(.026)
Δp_{t-1}	.324**	(.153)	.873***	(.163)		
Δp_{t-2}					-.588**	(.221)
V_{t-1}	.486***	(.105)				
V_{t-2}						
ΔGDP_{t-1}					3.48***	(.816)
ΔL_{t-1}			.787**	(.368)		
$p_{t-1}-p^e_{t-1}$	-.302***	(.080)	-.605***	(.159)		
Adj. R^2	.77		.59		.45	
SBC	-49.8		-37.7		-33.4	
			p-values		p-values	
J-B	.96		.09		.96	
LM(1)	.31		.26		.53	
LM(2)	.63		.02		.34	
LM(4)-heter	.85		.98		.40	
RMSE	.049		.018		.071	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. ***, **, * denote for statistical significance at the 10%, 5% and 1% level, respectively.

Appendix 2

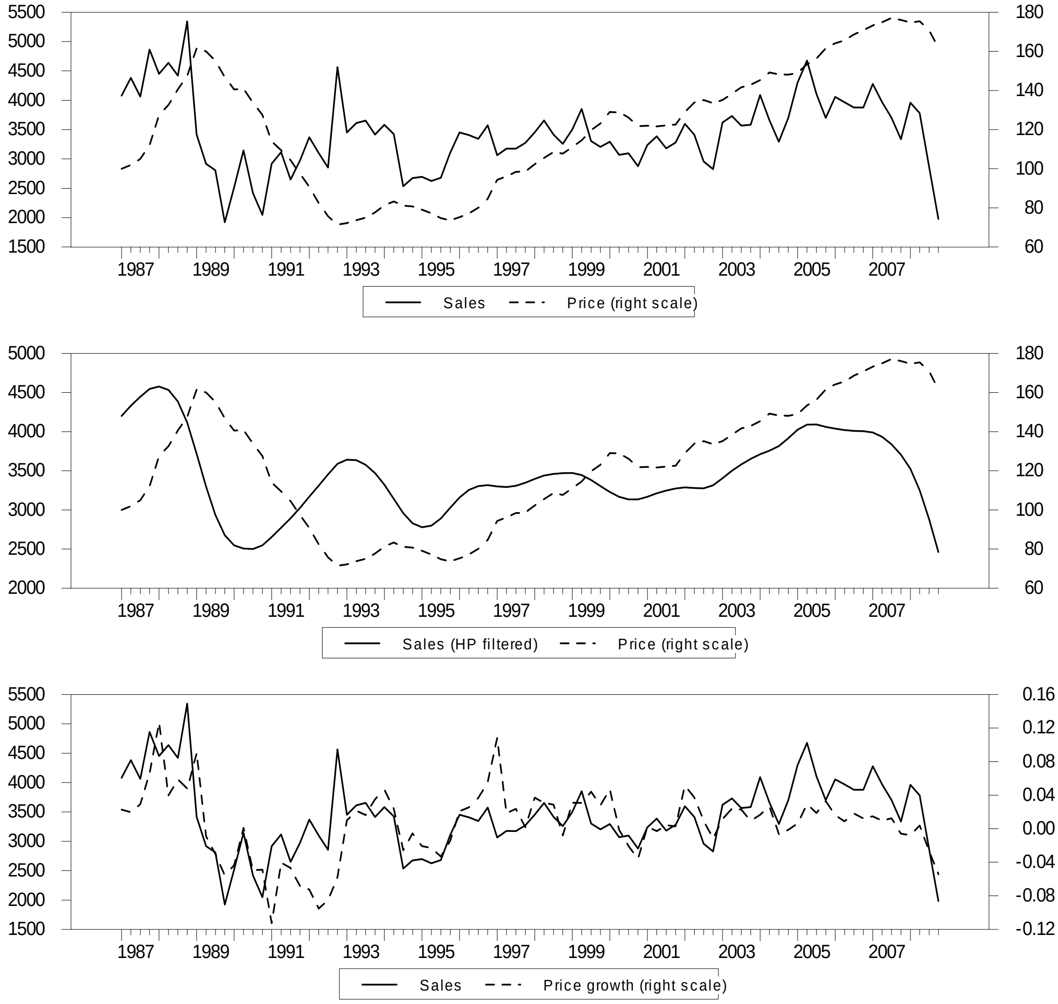

Figure A2 Real housing price index, sales volume and housing price growth over 1987-2008 in the Helsinki metropolitan area

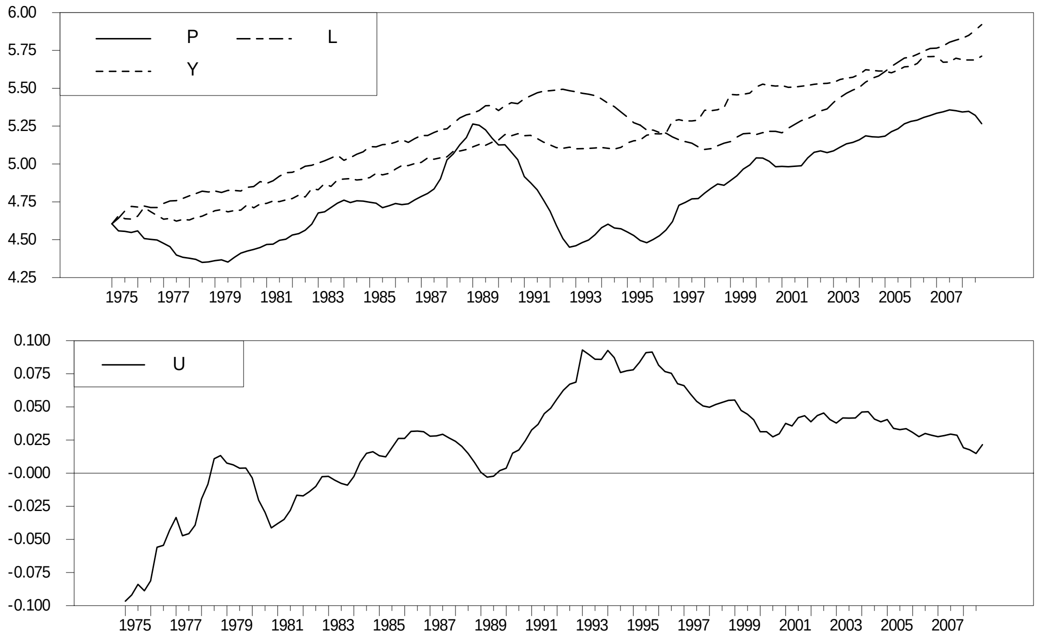

Figure A3 Real housing price index (P), real aggregate income index (Y), loan-to-GDP ratio index (L) and user cost of housing (U) over 1975-2008 in the Helsinki metropolitan area[22]

Table A6 Cointegration test statictics using HMA data

H_0 (rank)	Trace test value (p-value)
R = 0	36.3
R ≤ 1	4.0

P-value in the test for exclusion of *t* .42
P-value in the test for exclusion of *t* and *U* .49

Estimated long-run relation (standard errors in parenthesis):

$$P = 1.24 + \underset{(.132)}{.236} {*} Y + \underset{(.149)}{.451} {*} L$$

The tested model includes four stochastic variables (real housing price, sales volume, real aggregate income and lona-to-gdp ratio) and incorporates two lags in differences, three centered seasonal dummies and a deterministic trend in the long-run dynamics. 95% quantiles are 35.5 (R = 0) and 17.9 (R = 1). Exclusion of the variables is tested by the Bartlett small sample corrected LR test by Johansen (2000).

[22] L represents the nationwide loan-to-GDP ratio.

Table A7 P-values in the Granger non-causality tests using HMA data

Dependent Variable		Explanatory variable						
		ΔH	V	ΔY	ΔL	ΔU	eqe	Adj. R^2
	Δ Price (-.029)	**.00**	**.00**				**.02**	.71
	Volume (-.052)	.97	.00				.50	.46
	Δ Price (-.020)	**.00**	**.00**	**.06**	.59	.22	.32	.67
	Volume (-.076)	.66	**.00**	**.87**	.63	.51	.47	.47

The reported p-values are based on a standard F-test. Statistically significant values at the 10% level are bolded. The pairwise model includes two lags in differences while the six variable model includes one lag in differences. Both models include three seasonal dummies. Due to heteroscedasticity of the residuals in the model for ΔP, the p-values in the ΔP equation in the multiple variable model are based on a covariance matrix that is computed allowing for heteroscedasticity as in White (1980). HUOM! The Residuals in the equations for volume are non-normal…

Table A8 Prediction models for annual sales volume in HMA

	ECM1		ECM2		ECM3	
Constant	7.45***	(.949)	10.5***	(1.68)	10.7***	(1.50)
Trend	.015***	(.003)	.014***	(.004)	.014***	(.004)
Δp_{t-1}	-.330**	(.162)				
V_{t-1}	.177*	(.097)	.202	(.165)	.208	(.159)
V_{t-2}			-.345*	(.169)	-.371**	(140)
$p_{t-1}-p^e_{t-1}$	-.061	(.081)	-.032	(.106)		
Adj. R^2	.49		.52		.54	
SBC	-26.7		-27.9		-30.8	
			p-values		p-values	
J-B	.64		.86		.87	
LM(1)	.25		.09		.08	
LM(2)	.55		.22		.20	
LM(4)-heter	.04		.09		.10	
RMSE	.138		.130		.129	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. ***, **, * denote for statistical significance at the 10%, 5% and 1% level, respectively. Due to heteroscedasticity of the residuals, the covariance matrix in ECM1 is computed allowing for heteroscedasticity as in White (1980).

Table A9 Prediction models for annual price growth in HMA

	ECM1		ECM2		ECM3	
Constant	-2.48	(1.82)	-.008	(.021)	.001	(.023)
Δp_{t-1}	.521**	(.193)	.717***	(.171)	.667***	(.199)
Δp_{t-2}	.393	(.227)			-.344*	(.198)
V_{t-1}	.525***	(.154)				
V_{t-2}	-.265	(.158)				
$p_{t-1}-p^e_{t-1}$	-.387***	(.124)	-.298**	(.104)		
Adj. R^2	.65		.46		.33	
SBC	-33.7		-30.1		-25.5	
			p-values		p-values	
J-B	.27		.04		.75	
LM(1)	.09		.88		.72	
LM(2)	.18		.56		.21	
LM(4)-heter	.59		.89		.54	
RMSE	.026		.061		.046	

Standard deviation in parenthesis, J-B is the Jarque-Bera test for residual normality, LM(1) and LM(2) denote the Lagrange Multiplier test for the hypothesis of no first and second order residual autocorrelation and LM(4)-heter is the Lagrange Multiplier test for residual homoscedasticity. ***, **, * denote for statistical significance at the 10%, 5% and 1% level, respectively.

Aboa Centre for Economics (ACE) was founded in 1998 by the departments of economics at the Turku School of Economics, Åbo Akademi University and University of Turku. The aim of the Centre is to coordinate research and education related to economics in the three universities.

Contact information: Aboa Centre for Economics, Turku School of Economics, Rehtorinpellonkatu 3, 20500 Turku, Finland.

Aboa Centre for Economics (ACE) on Turun kolmen yliopiston vuonna 1998 perustama yhteistyöelin. Sen osapuolet ovat Turun kauppakorkeakoulun kansantaloustieteen oppiaine, Åbo Akademin nationalekonomi-oppiaine ja Turun yliopiston taloustieteen laitos. ACEn toiminta-ajatuksena on koordinoida kansantaloustieteen tutkimusta ja opetusta Turun kolmessa yliopistossa.

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