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ABSTRACT

In this study, regional differences in housing price dynamics are examined empirically using panel data models. We concentrate on examining the momentum dynamics and the reversion speed towards fundamental price level. The analysis can be seen as a test for the validity of conventionally used fixed-effects panel models to analyse regional housing price dynamics. Based on data over 1988-2012, the findings indicate that the regional differences are generally quite small in the Finnish market. We find a notable difference between Helsinki, by far the greatest city in Finland, and the other cities regarding the strength of the momentum effect, though. The results also provide evidence of cointegration between regional housing prices and income. The long-term coefficient on income considerably varies across cities.

JEL Classification: R21; R31; C33

Keywords: Housing price, dynamics, panel data, momentum, bubble

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Introduction

The understanding of housing price dynamics is of importance to a great number of agents: to portfolio investors, banks, real estate brokers and construction companies as well as to policy makers and most households. Unfortunately, the data that are used to study housing price dynamics empirically are generally problematic. One complication is the relatively low frequency of data on housing prices and on the fundamentals that significantly affect housing prices. Reliable housing price information is typically available only at the quarterly or, at best, monthly frequency. Moreover, data on some of the economic fundamentals that are expected to drive housing prices, such as income, are only available at an annual frequency in many markets.

Given the low frequency of data and the relatively short sample periods available, the econometric analyses examining housing price dynamics often exhibit small-sample complications. Therefore, a number of studies (e.g. Bramley and Leishman, 2005; Capozza et al., 2004; Harter-Dreiman, 2004; Hort, 1998; Lamont and Stein, 1999) investigate housing price dynamics using panel data models. By combining cross-sectional data, for instance multiple cities or countries, with the time series property of data, panel data analysis increases the number of observations and thereby the degrees of freedom in econometric modelling.

A potential complication with many of the panel data analyses is the use of ‘conventional’ fixed-effects models. ‘Conventional’ refers here to the fixed-effects models that assume housing price dynamics to be the same in every region included in the analysis (‘fixed main-effects models’). The parameters reported in the previous panel data analyses investigating housing price dynamics are generally average values across all the regions included in the data. However, there are reasons to believe that the housing price elasticities with respect to fundamentals substantially vary across distinct regional housing markets (Capozza et al., 2004; Davis and Heathcote, 2007). Similarly, the magnitude of autocorrelation in housing price movements, i.e., housing price ‘momentum’, as well as the speed of adjustment towards the long-run fundamental price level may notably differ between housing markets (Capozza et al., 2004; Glaeser et al., 2008). Therefore, the use of conventional fixed-effects panel data models may yield misleading results regarding regional housing price dynamics. This may lead to suboptimal policy and investment decisions being made.

This study aims to examine empirically the magnitude of regional differences in housing price dynamics using annual data for 14 Finnish cities for the period 1988-2012. In particular, it is investigated whether there are large regional differences in the momentum effect (‘bubble builder’) of housing prices, in the speed of adjustment of housing prices towards their long-run fundamental level (‘bubble burster’), and in the elasticity of housing prices with respect to income. The results of baseline model, i.e. the conventional fixed-effects panel model, are compared with those of less restrictive models where the dynamics are allowed to vary across cities. For this purpose, fixed interaction-effects panel models are estimated. As the aim is to test whether there are significant differences in the housing price dynamics across regions, the analysis can also be seen as a test for the validity of conventionally used fixed-effects models to analyse regional housing price dynamics. This appears to be the first panel data analysis formally testing for regional differences in the momentum and reversion dynamics of housing prices. Regarding the future research themes suggested by Bramley and Leishman (2005), the contribution of this article lies most prominently in modelling more explicitly the regional variations in housing price response behaviour.

The results indicate that the regional differences in short-run housing price dynamics across Finnish cities are relatively small: the analysis shows significant regional differences regarding neither the speed of adjustment towards the fundamental price level nor the coefficients on lagged fundamentals, with the exception of lagged housing stock change. However, the analysis presents evidence of notable regional variation in the momentum dynamics. This variation takes place between Helsinki, by far the greatest and most densely populated city in Finland, and the other cities.

In addition, the results provide evidence of cointegration between housing prices and per capita income. While cointegration is detected between housing prices and income in all the 14 cities, the long-run income elasticity considerably varies across cities. In general, the elasticity is the greater the larger is the city.

Overall, the findings propose that it can be misleading to rely on conventional fixed-effects panel models that do not allow for regional variation in the parameters other than the intercept. These models may lead to worse forecasts and investment and policy decisions than models that cater more rigorously for regional differences.

The study is organised as follows. The reasons to expect notable regional differences in housing price dynamics are discussed in the next section. Then, the empirical methodology is described, after which the data used in the empirical analysis are presented. The empirical results are reported in the fifth section. In the end, the study is summarized and concluded.

Considerations on regional differences in housing price dynamics

Land leverage and regional housing price dynamics

The value of land is expected to be a central factor causing regional differences in housing price elasticities with respect to economic fundamentals. Since desirable land is largely non-reproducible, changes in the demand for housing are likely to have substantial influences on the price of the land component of housing. By contrast, housing demand changes are expected to have a notably smaller impact on the real value of structures. A likely outcome is that the residential land value volatility is considerably greater than that of housing structures. Empirical evidence supporting this suggestion is provided in Davis and Heathcote (2007) and Gyourko and Saiz (2006).

The different dynamics of land prices and construction costs indicate that housing price dynamics should be quite different in regions where the value of housing is largely accounted for by the value of land, i.e, where the 'land leverage' is high, compared to regions where land's share of house value is relatively small, i.e., where the land leverage is low (Bostic et al., 2007; Bourassa et al., 2011; Davis and Heathcote, 2007).[1] Obviously, land leverage is closely related to supply constraints on

[1] Bourassa et al. (2009) show that housing price dynamics also can vary across individual units within a metropolitan area. The variation in dynamics across individual units is out of the scope of this paper.

land. Other things equal, tighter supply constraints and lower supply elasticity – due to zoning policies or other reasons increasing the scarcity of suitable residential land – lead to greater land leverage.

Conventionally, panel data analyses on housing price dynamics have been based on fixed-effect models that assume the parameters on fundamentals to be the same regardless of the region (e.g. Harter-Dreiman, 2004; Hort, 1998; Lamont and Stein, 1999; Jud and Winkler, 2002). Given that land leverage appears to substantially vary between regions (Bostic et al., 2007; Davis and Heathcote, 2007; Davis and Palumbo, 2008), the assumption of similar parameters across regions may be faulty. In the Finnish case, for instance, the effect of income growth on housing price level is expected to be greater in Helsinki, where land leverage is high, than in the considerably smaller city of Rovaniemi, where the value of land is considerably lower.

In the conventionally used fixed-effects models, housing appreciation rates are allowed to vary across areas not only due to regional differences in the evolution of market fundamentals but also because of location-specific fixed-effects. These fixed-effects represent the residuals of housing price appreciation attributable to location (Jud and Winkler, 2002). The important question is: what is the factor that the fixed-effects try to cater for? In other words, why would prices in one location rise more than in another location even if the development of the market fundamentals was exactly the same in both markets? Obviously, the reason for the different appreciation rates is the market specific elasticities of housing prices with respect to fundamentals.

The conventional fixed-effects models cannot cater properly for the regional heterogeneity of housing price elasticities. For instance, a fixed-effects model that assumes similar dynamics across regions would predict housing price growth to be greater in an area with a larger estimated intercept (fixed effect) both when housing demand is increasing and when the demand is decreasing. However, in an area with high elasticity of housing prices with respect to economic fundamentals, prices should decline faster when the demand is decreasing. That is, the different average appreciation rates across regions are generally due to the different development of the fundamentals *and* due to the varying housing price elasticities across markets. A model that does not allow for regional variation in the dynamics may yield biased predictions for regional housing price development and a biased picture concerning regional housing price dynamics in general.

Housing price momentum and reversion towards the fundamental price level

Housing prices have been shown to exhibit notable short-run persistence and long-run mean reversion (e.g. Beracha and Skiba, 2010; Capozza et al., 2004; Case and Shiller, 1989, 1990; Malpezzi, 1999; Roed Larsen and Weum, 2008). Abraham and Hendershott (1996) call the short-term positive autocorrelation, i.e., the short-term momentum, a ‘bubble builder’ due to its tendency to often drive housing prices further away from their fundamental level. The reverting tendency of housing prices towards their long-run fundamental level, in turn, is often referred to as a ‘bubble burster’. The magnitude of momentum and strength of reversion towards the fundamental level are of great importance regarding housing price dynamics, not least because of their predictability implications and their influence on the occurrence and magnitude of housing price cycles and bubbles. Importantly, the momentum and reversion dynamics may significantly vary across regions.

In the housing market, backward-looking expectations are likely to strengthen the momentum effect. For instance Capozza et al. (2004), Dusansky and Koc (2007) and Fraser et al. (2008) present evidence of backward-looking expectations in the housing market. Various informational factors can influence the significance of such feedback effects in a given market.

Clapp et al. (1995) suggest that higher population density should foster more, better and prompter information concerning the housing market, since information production is subject to positive scale economies. Moreover, in markets with a greater number of transactions, information costs are lower and, therefore, prices should respond more rapidly to changing fundamentals (Capozza et al., 2004). In line with these arguments, empirical evidence suggests that people show stronger behavioral biases driven by psychological constraints when the asset is harder to value (Hirshleifer et al., 2013; Kumar, 2009). Generally, housing is easier to value in an area with a greater number of transactions and thereby greater information flows.

Overall, these informational factors suggest that in larger and more densely populated metropolitan areas with more liquid housing market, behavioral biases such as backward-looking expectations should be less significant and housing demand should more rapidly adjust to shocks. Therefore, in these kinds of regions the adjustment towards fundamental price level should be more rapid and the momentum effect is expected to be weaker.

However, there can also be other reasons than the informational factors that cause regional variation in housing price momentum and reversion. Capozza et al. (2004) hypothesise that higher real construction costs are correlated with slower reversion and greater serial correlation. Construction costs vary between regions because of material and labour costs and also due to unpriced supply restrictions. The results of Hwang and Quigley (2006) emphasize the importance of local regulation (i.e., administrational supply restrictions) on housing market dynamics. Furthermore, in less densely populated areas supply may be able to adjust more rapidly than in areas with greater density and scarcity of land (Glaeser et al., 2008).

Since the informational and (other) structural factors may have opposite effects on housing price persistence and reversion dynamics at the market level, it is essentially an empirical question to study the variation of price dynamics across regional housing markets. Unsurprisingly, some previous studies report notable regional differences in housing price dynamics (e.g. Abraham and Hendershott, 1996; Bramley and Leishman, 2005; Capozza et al., 2004; Holly et al., 2010; Hwang and Quigley, 2006; Malpezzi, 1999). On the other hand, Englund and Ioannides (1997) find the autocorrelation structures to be strikingly similar across countries.

The panel data analyses studying housing price dynamics typically have not considered potential regional differences. There are some exceptions, however. Holly et al. (2010) allow the dynamics to vary across U.S. regions, and in Malpezzi (1999) the adjustment speed towards fundamental price level varies across U.S. areas with varying degrees of land use restrictions. In addition, Abraham and Hendershott (1996) estimate separate fixed-effects models for coastal cities and inland cities in the U.S. Bramley and Leishman (2005), in turn, divide their U.K. panel dataset into three broad area types to allow for regional variation. They find that housing market dynamics differ between markets that are characterised by low demand and those under high demand pressure. Despite some recent research in the topic, the statement of Bramley and Leishman (2005) still holds today: “There are clearly considerable opportunities to build further on this in future research. The model could address the dynamics of change and regional interactions in a more sophisticated way and could more explicitly build in regional variations in response behaviour.”

The aim of this study is to bring new empirical evidence on the theme by estimating panel data models that allow for variation in regional housing price dynamics. Unlike in previous studies, the significance of regional variation in both the momentum and equilibrium-adjustment parameters is tested formally. Regarding the future research themes suggested by Bramley and Leishman (2005), the contribution of this article lies most prominently in modelling more explicitly the regional variations in housing price response behaviour.

Empirical model and methodology

For the purposes of this study, an empirical model that corresponds well to the actual city level housing price dynamics is needed. Based on the findings reported in the earlier literature, this model should capture 1) the short-run momentum in housing prices, 2) the long-run tendency of housing prices to revert towards fundamental price level (‘fundamental level’ and ‘long-run equilibrium level’ are used as synonyms in this study), and 3) the potential short-run effects of market fundamentals on housing price growth. The baseline fixed-effects panel data model (1) is closely related to those used by e.g. Lamont and Stein (1999) and Malpezzi (1999):

$$\Delta p_{it} = b_{0i} + b_1\Delta y_{it-1} + b_2\Delta d_{it-1} + b_3\Delta s_{it-1} + b_4\Delta ir_{t-1} + b_5\Delta p_{it-1} + b_6(p - p^*)_{it-1} + u_{it} \quad (1)$$

The model captures the aforementioned features in a simple error-correction framework. In (1), the dependent variable is real housing price change (Δp) in city i in period t, while the explanatory variables include the one period lagged deviation of housing prices from their long-run equilibrium level ($p - p^*$), the previous period housing price growth, and lagged changes in four fundamentals that may affect the short-run housing price dynamics based on the life-cycle model of the housing market (see e.g. Meen, 2001). These fundamentals include the real per capita income (y), population (d), housing stock (s), and the real after-tax interest rate (ir). In contrast with the Lamont and Stein (1999) and Malpezzi (1999) models, (1) includes only lagged fundamentals. This is because the aim is to estimate models that could be used for prediction purposes. Moreover, the fundamentals are potentially endogenous with respect to housing prices. The model also includes a deterministic constant (b_0) that is allowed to vary across cities. Since we use a nationwide measure for ir, we have dropped i from this variable.

A long-run equilibrium price level can be defined as one from which there is no systematic tendency to depart. Following Harter-Dreiman (2004), Holly et al. (2010), Lamont and Stein (1999) and Malpezzi (1999), among others, the long-run equilibrium housing price level in city *i* during period *t* is computed as the long-term relationship between real housing price level and real per capita disposable income:

$$p_{it}^* = \delta_i^* \, y_{it} \tag{2}$$

The long-term coefficient on income (δ^*) is allowed to vary across cities and differ from one, since the equilibrium price-to-income ratio is not necessarily constant over time and space. The equilibrium ratio and its temporal evolution are dependent on the elasticities of supply of labor and housing, and on the driving forces behind metropolitan population growth (DiPasquale and Wheaton, 1996; Malpezzi, 1999). Obviously, $\delta^e > 0$.

As p and y are both non-stationary in levels but stationary in first differences (see the data section), we use the Johansen Trace test to test for cointegration between the variables. The cointegration analysis is conducted and the long-run equilibrium relationship is computed separately for each city, and it works as a specification check for (2). If the relationship given in (2) is found to be stationary, i.e., p and y are pairwise cointegrated, and if the parameter δ^* is found to be stable over time, then the relation can be regarded as one towards which housing prices tend to adjust and from which the price level cannot drift away in the long run. The price level can temporarily deviate from p_t^*, though. Instead, if cointegration could not be detected between housing prices and income, the relation could not be regarded as a long-run equilibrium relation for housing prices.

In addition to the conventional fixed-effects model (i.e. the baseline model show in (1)), we estimate fixed interaction-effects models that allow for different parameter estimates across cities or city groups:[2]

$$p_{it} = b_{0i} + b_{1i}\Delta y_{it-1} + b_{2i}\Delta d_{it-1} + b_{3i}\Delta s_{it-1} + b_{4i}\Delta ir_{t-1} + b_{5i}\Delta p_{it-1} + b_{6i}(p - p^*)_{it-1} + u_{it} \quad (3)$$

$$p_{it} = b_{0i} + b_{1j}\Delta y_{it-1} + b_{2j}\Delta d_{it-1} + b_{3j}\Delta s_{it-1} + b_{4j}\Delta ir_{t-1} + b_{5j}\Delta p_{it-1} + b_{6j}(p - p^*)_{it-1} + u_{it} \quad (4)$$

Specification (3) includes interaction dummies that allow the impacts of explanatory variables to vary across all 14 cities. Specification (4), in turn, is estimated using two different regional divisions. In the first one, the coefficients vary between Helsinki, by far the greatest and most supply restricted city in Finland, and the other cities on aggregate (i.e., the coefficient estimates are the same for all cities outside Helsinki – the intercept varies between all the cities in all the estimated models, though). The second division makes distinction between three groups of cities: we estimate different coefficients for Helsinki, other growth centres, and for the smaller and more peripheral cities. In other words, j includes either two groups (Helsinki and other cities) or three groups (Helsinki, other growth centres, and all the other cities).

The sign of the speed of adjustment parameters (b_6) is expected to be negative. That is, when the price level is higher than its long-run fundamental level, the price growth rate is expected to be small so that housing prices adjust towards their equilibrium

[2] An alternative to using fixed-effects models with interaction terms to allow for regional differences would be to estimate random-effects models. The random-effects models are not estimable with the data used here. This is most likely because of the small variance of random effects. This is a well-known complication in random-effects models.

level with income. Furthermore, given the notable frictions in the housing market and the potential feedback effects (backward-looking expectations), the momentum coefficient (b_5) is anticipated to be positive – fast housing price growth in this period predicts fast growth in the next period.

If the lagged fundamentals significantly affect the dynamics of Δp, the parameter signs are expected to be positive on Δy and Δd, and negative on Δir and Δs. If housing price dynamics are dominated by the reversion towards long-run equilibrium level and the short-run momentum, the coefficients on the lagged fundamentals may well be insignificantly different from zero. Therefore, we also formally test for the exclusion of the lagged fundamentals using the Likelihood Ratio (LR) test. However, of main interest are LR tests on the hypothesis that the momentum and reversion parameters are the same across cities.

A complication regarding a dynamic panel data model that includes the lagged dependent variable is endogeneity bias (Wooldridge, 2002). Therefore, the panel models are estimated using the Arellano-Bover/Blundell-Bond Generalized Method of Moments (GMM) technique. The two-period lagged transaction volume (total number of transactions) is used as the instrument for one-period lagged housing price change. The two-period lagged volume is used instead of the one-period lagged one, since transaction volume leads housing price movements. This is in line with previous empirical evidence for the Finnish market (see Oikarinen, 2012). The correlation between the instrument and the price change varies between 0.54 – 0.83 across cities, and is 0.72 on average.

Furthermore, the estimated models may include spatial autocorrelation, i.e., the dependent variable may be jointly determined with the other cities' development and its own characteristics (Elhorst, 2010). Indeed, previous research has found evidence of spatial autocorrelation across housing markets (Holly et al., 2010; Kuethe and Pede, 2011). Therefore, the Moran's I test is conducted to investigate whether the models exhibit spatial autocorrelation. We use several different 'distance' measures in the tests, including the number of inhabitants, distance from Helsinki and various industry shares (as a proxy for the similarity of regions' economic bases). Regardless of the distance measure, the test statistics do not show evidence of significant spatial autocorrelation.

Data

The empirical analysis is based on annual data for 14 Finnish cities for the period 1988-2012. The cities include the ten largest ones in Finland and some other regional centres. It is reasonable to use data only since 1988, as the financial market deregulation that took place in Finland during the late 1980s induced a structural break in housing price dynamics (Oikarinen, 2009a, 2009b). The data are sourced from the Statistics Finland database unless mentioned otherwise.

The hedonic housing price indices computed by Statistics Finland measure the housing price development. The indices are based on transactions of privately financed apartments in the secondary market. There are good reasons to focus on the privately financed sector: In Finland, privately financed housing can be bought and sold at market prices without any restrictions, whereas selling prices and rental prices are controlled in the publicly regulated (i.e. subsidized) sector. Furthermore, the data consist only of apartments, since data on apartments are more reliable than data on the other housing types: apartments are a substantially more homogenous group than the other housing types, and a notably greater number of transactions take place in the apartment market than in the market for other housing types. That is, the use of apartment data diminishes the heterogeneity problem that is associated with housing price data even when hedonic indices are employed.

The city-level income per capita variable (*y*) is the aggregate after-tax household income divided by the population of the city. Since the mortgage interest rates are practically the same all over the country, we use the nationwide after-tax mortgage rate to measure *ir*. The mortgage rate data are provided by the Bank of Finland. Finally, *d* is measured as city-level end-of-year population, and *s* as the end-of-year total housing stock (m^2).

The panel models are estimated using real variables. Thus, *p*, *y*, and *ir* have been deflated by the cost of living index. Natural logs of all the variables except for the interest rate are used. Table 1 presents summary statistics on housing price growth. In Table 1, “Group” refers to the city group in which each of the cities are allocated. Group 1 only contains Helsinki, the other growth centres form Group 2, and Group 3 consists of the rest of the cities. The cities in Group 2 have grown more rapidly and are larger than those in Group 3, and Group 3 cities can be generally considered to be less supply constraint and to have lower land leverage than Group 2 cities. Nevertheless, the most notable difference in land leverage and supply restrictions is that between Helsinki and the other cities. Note also that there have been several changes in the administrational geographic boundaries of the cities after 2008. The data correspond to the city boundaries prior to those changes.

[Table 1 around here]

All the housing price series are nonstationary based on the Augmented Dickey-Fuller (ADF) unit root test, while housing price growth is stationary (Table 1). This is in line with previous empirical evidence reported for numerous housing markets. The unit root tests are applied individually for each city, since the cointegration tests are conducted separately for each city. All the other variables are difference stationary as well (these ADF test are available from the authors upon request). Table 1 also reveals that housing price growth is highly positively autocorrelated in all of the cities.

Empirical results

Before estimating and comparing the panel models, the cointegration analysis is conducted. The Johansen Trace test results suggest that p and y are cointegrated in all the regions included in the analysis (Table 2). This is in line with the results by Malpezzi (1999) which present evidence of cointegration between metropolitan housing prices and per capita income in the U.S. The cointegrating relations estimated by the Johansen Maximum Likelihood methodology are sensible with respect to the size of coefficients on y, and the stability of these relations cannot be rejected based on the recursive Max test (Juselius, 2006). Therefore, the long-term relations between *p* and *y* can be considered as reasonable measures for long-term equilibrium housing price levels in the cities.[3]

[Table 2 around here]

The coefficient on *y* varies significantly across regions. In line with Malpezzi's (1999) findings for the U.S., the income elasticity of housing prices is greater in larger cities that are generally more supply restricted due to the scarcity of vacant land in attractive sites. The coefficient is the greatest in the three biggest cities, and notably greater in Helsinki – by far the largest city – than in the second and third largest cities. Moreover, the income elasticity of housing prices is the lowest in the smallest city (Kajaani). The simple correlation between the coefficient on *y* and population is as large as 0.83, while that between the coefficient on *y* and the price elasticity of housing supply (reported in Oikarinen et al. (2014)) is –0.88. These correlations are as expected, since land leverage is generally smaller in the smaller and less supply restricted cities. The land supply restrictions and land leverage are not the only factors affecting the coefficient, though (DiPasquale and Wheaton, 1996).

Figure 1 present the evolution of real housing prices and their long-run fundamental level in the cities during 1988-2012. The graphs show clearly the price overshot in the

[3] The low power of the Johansen Trace test in small samples is well known. Therefore, some authors use panel data cointegration tests that have greater power properties. Since the aim is to allow for regional variation in the coefficient on *y* and to conduct a recursive analysis on the stability of the estimated relation for each city separately, the Trace test is used instead of a panel cointegration test. Moreover, small-sample corrected Trace values are reported, and the relatively low power of the Trace test is not a problem here, since the hypothesis of no cointegration is rejected for all the cities.

late 1980s that followed the abolishment of financial market deregulation in Finland. After the price overshot, the dramatic drop in housing prices was strengthened by the deep recession in the Finnish economy during the early and mid 1990s.

[Figure 1 around here]

The deviation of observed housing price level from the cointegrating relationship is used in the error-correction term in the panel models. The dynamic GMM panel estimation results are presented in Table 3. All reported models include the lagged fundamentals, since the fundamentals cannot be excluded based on the LR test. The Sargan test accepts the hypothesis of exogeneity of instruments in all the models, and the Moran's I test does not show evidence of spatial autocorrelation in any of the specifications. Hence, these models appear to be reasonably well specified.

[Table 3 around here]

The estimated parameters generally have the expected sign, are of sensible magnitude, and are statistically significant. An exception is Interaction model 1, where several parameters are not statistically significant and some have an unexpected sign. The positive coefficients on one-year lagged interest rate change also might be regarded as unexpected. The positive sign could imply that there is an initial overreaction to an interest rate chance and prices get 'back in line' a bit later. Anyhow, we focus on investigating the momentum and reversion parameters – their variation across regions, in particular.

In the baseline model, in which all the coefficients on stochastic variables are restricted to be the same across cities, the coefficient estimate 0.6 on $\Delta p_{i,t-1}$ indicates considerable momentum in housing prices: faster price growth this year predicts more rapid growth next year. Both the momentum parameter and the adjustment speed towards the long-run relation, –0.4, are somewhat greater (in absolute value) than those reported based on similar type of models by Lamont and Stein (1999) and Holly et al. (2010) for the U.S.

In Interaction model 1, where all the coefficients are allowed to differ across all the 14 cities, the estimated coefficients notably vary: the momentum parameter from 0.1 to 1.0 and the 'bubble burster' parameter from 0.1 to –0.8. This variation is not statistically significant, though, as the p-values in the LR test are 0.17 and 0.51, respectively. Since Helsinki notably differs from the other cities, but the others cities generally form a relatively uniform group, the inability to reject the similarity of momentum and reversion parameters across cities in Interaction model 1 may be due to the relatively small coefficient variation between the 13 cities outside Helsinki. Therefore, we estimate models that make a distinction between Helsinki and the other cities (Interaction models 2 and 3). We also estimate models that allow for differences between Helsinki, the other growth centers and the remaining cities. However, as the differences between the other growth centres and the rest of the cities are only slight and statistically clearly insignificant, we do not report those results.

In Interaction models 2 and 3, the momentum parameter is a borderline case regarding statistically significant variation between Helsinki and the other cities. The point estimate is approximately 0.4 for Helsinki and 0.6 for other cities. The estimated weaker momentum in Helsinki is in line with the suggestions of Clapp et al. (1995) according to which higher population density should foster more, better and prompter information concerning the housing market, and of Hirshleifer et al. (2013) and Kumar (2009) based on which people show stronger behavioral biases (such as the feedback effect) when the asset is harder to value. The finding further implies that the informational factors have a more significant impact on momentum dynamics than the supply side elasticities do. While the difference in the point estimates is statistically significant at the 6% level and can be regarded as being of economic significance (see Figure 2 and the discussion below), it should be noted that it is not statistically significant at the 5% level.

In Interaction model 2, the adjustment speed towards long-run relation is somewhat slower in Helsinki (30% per year) than elsewhere (40%). Even this kind of variation could have considerable impacts regarding price predictions: if regional housing prices were for instance 20% below their fundamental level, the model would predict a 2%-point smaller housing price growth in Helsinki than elsewhere during the next year, and slower growth in Helsinki during several years from the second year onwards as well. The difference between these 'bubble burster' parameters is not statistically significant, though, and Interaction model 3 in which the speed of adjustment parameter does not vary across regions is preferred over the other models based on the adjusted R^2.

Note also that the only lagged fundamental variable for which the analysis indicates significant variation between Helsinki and the other cities is housing supply.

Therefore, the coefficients on the other fundamentals are restricted to be the same across cities in Interaction models 2 and 3.

The price dynamics are investigated in more detail in Figure 2. In the figure, the baseline adjustment paths of housing prices are computed based on Interaction model 3. To illustrate the influence of potential differences in the reversion parameter on the price dynamics (this is relevant here, as substantial and statistically significant differences can exist within many other countries), Figure 2 also shows the adjustment paths based on Interaction model 2. In addition to showing how the momentum effect and reversion towards the fundamental price level influence the dynamics, Figure 2 gives insight regarding the economic significance of the differences between Helsinki and other cities.

[Figure 2 around here]

The two graphs in the upper part of Figure 2 assume an income change that increases the equilibrium housing price level by one percent, i.e., an income growth of approximately 0.5% in Helsinki and between 1.3% and 0.7% for the other cities. The graphs in the mid part show the housing price evolution after 1% increase in the housing price level that is not accompanied with a change in the income. Finally, the low part of Figure 2 shows the response of housing prices to 1% increase in income. The graphs in the left hand side show the ‘baseline’ dynamics based on Interaction model 3. The right hand side graphs, in turn, show the dynamics based on Interaction model 2, i.e., as if the speed of adjustment differed between Helsinki and the other cities.

The graphed reaction patterns show that the adjustment path of housing prices is more cyclical and housing prices are more prone to notable overshots in the other cities than in Helsinki. This is due to the stronger momentum in the other cities. The difference between the baseline graphs and those computed from Interaction model 2 is only slight. For the other cities, the parameter estimates are practically the same regardless of whether Interaction model 2 or 3 is used. However, for Helsinki there is a more considerable difference regarding the adjustment speed towards long-term equilibrium. The faster adjustment speed in Helsinki assumed in the left-hand side graphs induces slightly more cyclical adjustment path: while prices tend to revert faster towards their long-term fundamental level, the faster reversion induces greater price changes thereby causing a greater eventual price overshot. As mentioned, the observed difference is only small and hypothetical as Interaction model 2 is preferred by the model statistics.

Interestingly, the tendency of housing prices to oscillate around the long-run fundamental level seems to be most prominent in the two least densely populated cities of our sample, i.e., in Kajaani and Rovaniemi (see Figure 1). This is in line with Clapp’s (1995) arguments. Nevertheless, the greater tendency of housing prices to oscillate around the long-run relationship does not necessarily imply that prices are expected to be more volatile in the smaller cities, at least in the long run. The low part of Figure 2 illustrates that income changes of similar magnitude cause larger housing price movements in Helsinki, where the long-run income elasticity is notably greater than that in the other cities.

Note that the reaction patterns shown in Figure 2 resemble those of ‘Region III’ derived in Capozza et al. (2004). The empirical evidence by Capozza et al. (2004) shows that oscillatory convergence generally characterizes housing price dynamics in the U.S. metro areas too.

Conclusions

The aim of this study is to examine empirically the magnitude of regional differences in housing price dynamics using panel data models that allow for variation in the dynamics across housing markets. In particular, the aim is to investigate whether there are large regional differences in the momentum effect (‘bubble builder’) of housing prices and in the speed of adjustment towards the long-run fundamental level (‘bubble burster’). The analysis can also be seen as a test for the validity of conventionally used fixed-effects panel data models which assume the housing price dynamics to be the same across all the housing markets included in the model.

Applying the Arellano-Bover/Blundell-Bond GMM technique for a panel of 14 Finnish cities, the regional differences in short-run housing price dynamics are found to be relatively small. The results show statistically significant differences across Finnish cities regarding neither the speed of adjustment parameter towards the long-run fundamental price level nor the coefficients on lagged fundamentals, with the exception of the lagged housing stock change. However, the analysis indicates regional variation that is of economic significance in the momentum parameter, i.e., in the parameter on instrumented previous price change. This variation between Helsinki, by far the greatest and most densely populated city in Finland, and the other cities is a borderline case with respect to statistical significance, as the difference is significant at the 6% level but not at the 5% level. The estimated models are in line with a hypothesis according to which the informational factors have a more significant impact on short-term momentum dynamics than the supply restrictions do.

The analysis also provides support for cointegration between housing prices and income. While housing prices and income are cointegrated in all the cities, the long-run coefficient on income significantly varies across regions. In general, the coefficient is the greater the larger and more supply restricted is the city, as expected. While the greater momentum effect causes housing prices to be more prone to notable overshots in the smaller cities outside Helsinki, the notably greater long-run income elasticity of housing prices in Helsinki indicates that housing price volatility can be substantially greater in Helsinki than in the other cities, at least in the long horizon. Catering for both momentum and income elasticity issues, the findings indicate that the concentration of population to the largest centre does not *necessarily* cause greater overall housing price volatility in a country. This is due to the informational factors: information production is subject to positive scale economies.

This empirical analysis is conducted using data for a country that is relatively small in size and coherent in terms of culture and income. In a geographically larger and culturally and economically more diverse country than Finland, the regional differences are likely to be more pronounced. That is, in many cases the use of conventional fixed-effects models, in which only the intercepts are allowed to vary across regions, is likely to yield misleading conclusions regarding regional housing markets. Therefore, further empirical examination on regional differences in the housing price dynamics using data for other countries is desirable.

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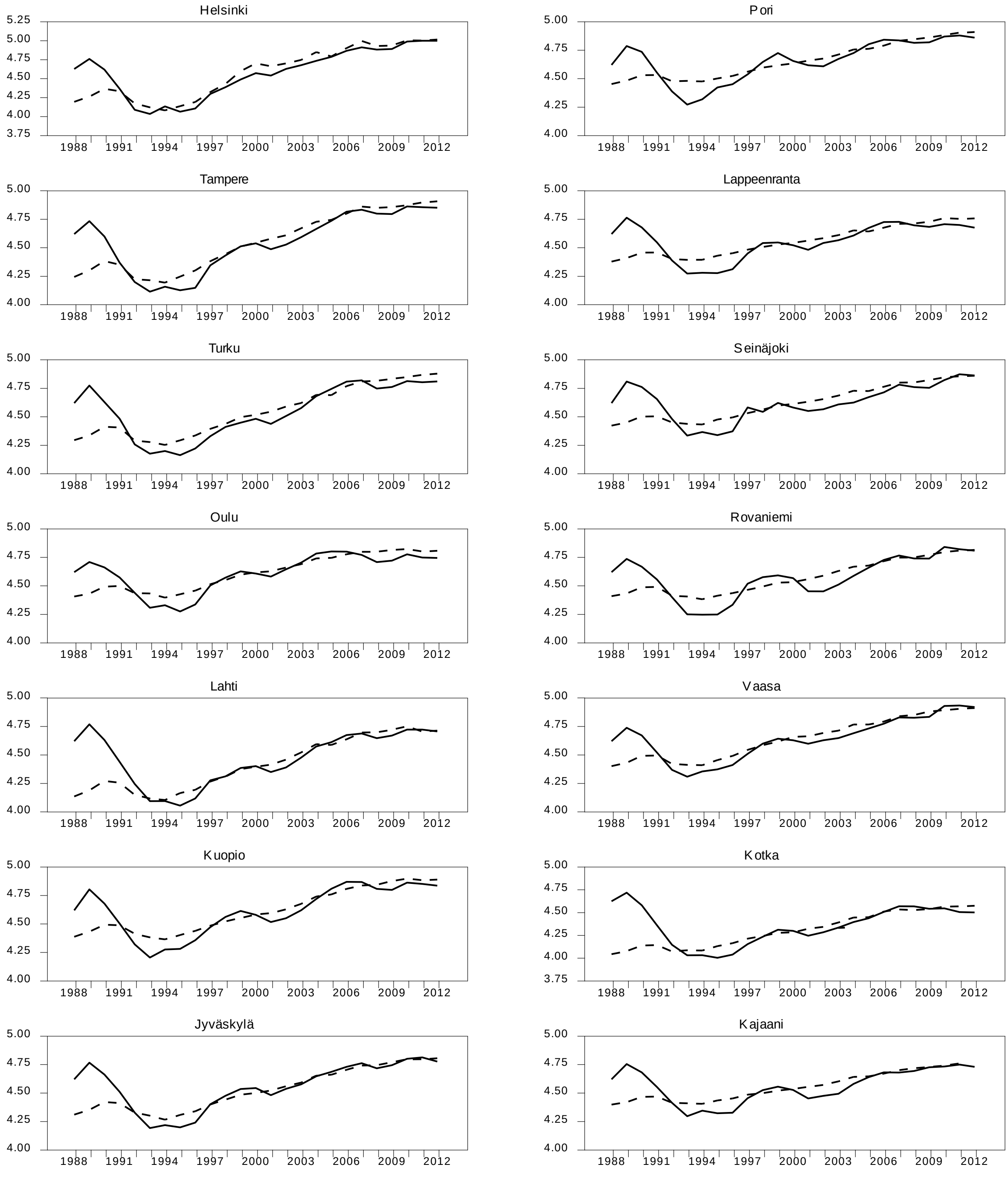

Figure 1. Real housing price indices and the long-run fundamental housing price level

The Figure shows the real housing price index and the equilibrium price level indicated by the cointegrating relation between housing prices and income ("long-run relation") in each of the 14 cities.

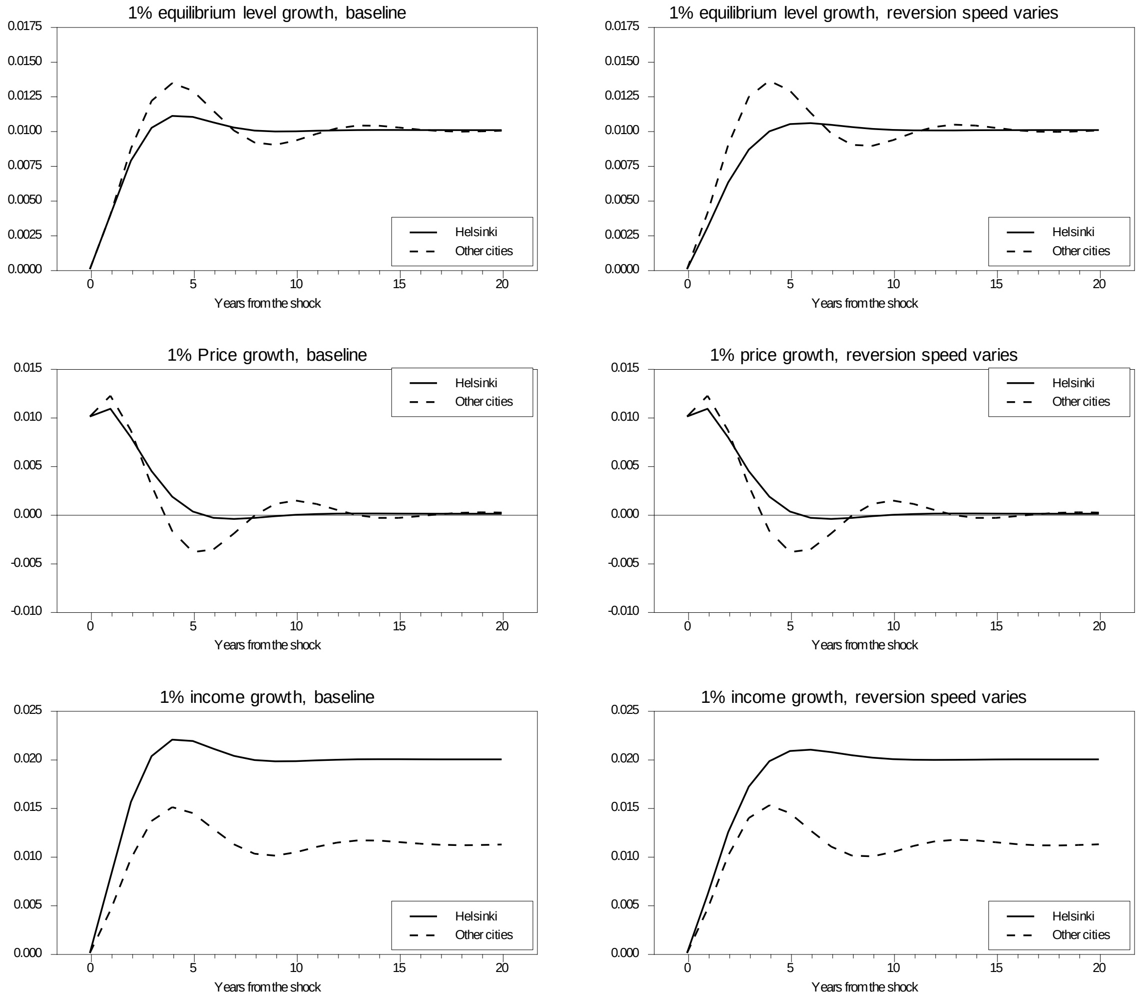

Figure 2. Housing price adjustment paths

The Figure shows the effects of various changes on the housing price behavior in Helsinki and in the other cities. "Baseline" refers to Interaction model 3, while "reversion speed varies" refers to Interaction model 2 where both the momentum parameter and the speed of adjustment towards the long-run cointegrating relation differ. "1% equilibrium level growth" stands for an income change that increases the long-run equilibrium level for housing prices by one percent. For the other cities, the long-term income elasticity of housing prices is computed as the average of the elasticities reported in Table 2.

Table 1. Descriptive statistics of real housing price growth

CITY	*Group*	*Mean*	*S.D.*	*1st Order Autocorrelation*	*ADF level (lags)*	*ADF difference (lags)*
Helsinki	1	0.015	0.112	0.478*	–1.01 (3)	–2.89** (0)
Tampere	2	0.010	0.094	0.523**	–1.63 (1)	–2.74** (0)
Turku	2	0.008	0.090	0.435*	–1.63 (1)	–3.21** (0)
Oulu	2	0.005	0.072	0.463**	–1.73 (1)	–2.49* (0)
Lahti	2	0.004	0.095	0.505**	–2.34 (1)	–2.91** (0)
Kuopio	2	0.009	0.093	0.445*	–1.93 (1)	–3.27** (0)
Jyväskylä	2	0.007	0.086	0.461*	–2.06 (1)	–3.08** (0)
Pori	3	0.010	0.085	0.481*	–2.12 (1)	–3.91** (0)
Lappeeranta	3	0.002	0.075	0.482*	–2.73 (1)	–3.10** (0)
Rovaniemi	3	0.008	0.085	0.506**	–2.26 (1)	–5.02** (3)
Vaasa	3	0.012	0.068	0.497*	–1.42 (1)	–2.95** (0)
Seinäjoki	3	0.010	0.087	0.196	–0.89 (0)	–4.32** (0)
Kotka	3	–0.005	0.090	0.638**	–1.45 (2)	–2.29* (0)
Kajaani	3	0.005	0.072	0.402*	–2.32 (1)	–3.38** (0)

The cities are ordered by population in 2008. * and ** denote statistical significance at the 5% and 1% level, respectively. Mean is the average log change of real housing price level, and S.D is the standard deviation of the log price change. In the Augmented Dickey-Fuller (ADF) unit root test, the critical values at the 5% and 1% significance levels are –2.99 and –3.75, respectively, in the test for levels (i.e., when an intercept is included in the test) and –1.96 and –2.67 in the test for differences (i.e., when the test does not include any deterministic variables). The number of lags included in the ADF tests is selected by the Schwarz Information Criteria.

Table 2. Cointegration test results

City	Trace test statics	p-value	Coefficient on *y* (standard error)
Helsinki	27.1**	.00	1.99 (.075)
Tampere	29.7**	.00	1.53 (.089)
Turku	15.2*	.05	1.42 (.098)
Oulu	22.5**	.00	.934 (.080)
Lahti	32.7**	.00	1.46 (.113)
Kuopio	35.6**	.00	1.09 (.055)
Jyväskylä	31.3**	.00	1.19 (.078)
Pori	30.5**	.00	.905 (.065)
Lappeeranta	27.2**	.00	.788 (.083)
Rovaniemi	21.8**	.00	.962 (.106)
Seinäjoki	16.5*	.03	.822 (.134)
Vaasa	29.2**	.00	1.09 (.057)
Kotka	34.9**	.00	1.19 (.115)
Kajaani	22.9**	.00	.781 (.098)

The Table shows the Johansen Trace statistics on hypothesis r = 0 (i.e., no cointegration). The reported Trace test statistics are small-sample corrected as suggested by Johansen (2002). * and ** denote statistical significance at the 5% and 1% level, respectively. p-value shows the probability value of the corresponding Trace test value, and the coefficient on *y* shows the estimated δ^e_i in the relationship $p_{it}* = \delta_i^e\, y_{it}$. The lag length in a test is selected by the Schwarz Information Criteria, and is one in differences in all the tests except in those for Helsinki, Turku, and Vaasa, in which there are two lags. The cities are ordered by population in 2008: the biggest city is at the top and the smallest city at the bottom. The residuals in the tested models are normally distributed based on the Jarque-Bera test, and non-autocorrelated based on the LM(1) test. The stability of the cointegrating relations over the sample period cannot be rejected based on the recursive Max-test (Juselius, 2006).

Table 3. Dynamic GMM panel model statistics

	Baseline model	Interaction model 1	Interaction model 2	Interaction model 3
$\Delta p_{i,t-1}$	0.580** (0.060)	0.065 – 1.03		
$(p_i - p_i^*)_{t-1}$	–0.360** (0.053)	(0.076) – (–.776)		–0.391** (0.061)
$\Delta p_{ij,t-1}$ Helsinki			0.379** (0.129)	0.387** (0.129)
$\Delta p_{i,t-1}$ Other cities			0.617** (0.067)	0.611** (0.061)
$(p_i - p_i^*)_{t-1}$ Helsinki			–0.300 (0.165)	
$(p_i - p_i^*)_{t-1}$ Other cities			–0.404** (0.065)	
$\Delta y_{i,t-1}$	0.287 (0.156)	–0.553 – 1.20	0.306 (0.161)	0.301 (0.160)
$\Delta d_{i,t-1}$	3.28** (0.923)	–5.06 – 8.59	3.21** (1.03)	3.24** (1.03)
$\Delta ir_{i,t-1}$	0.007** (0.003)	–0.010 – 0.025	0.008** (0.003)	0.008** (0.003)
$\Delta s_{i,t-1}$	–2.31** (0.479)	–10.6 – 0.956		
$\Delta s_{i,t-1}$ Helsinki			–9.69** (3.18)	–9.20** (3.07)
$\Delta s_{i,t-1}$ Other cities			–2.38** (0.571)	–2.44** (0.563)
R^2	.479	.578	.489	.488
Adj. R^2	.449	.416	.454	.456
LR, $\Delta p_{i,t-1}$		.17	.08	.06
LR, $p_i - p_i^*$		.51	.55	

The dependent variable is the log real housing price change, $\Delta p_{i,t}$. The two-year lagged transaction volume is used as an instrument for $\Delta p_{i,t-1}$. Standard errors are reported in the parenthesis. * and ** denote statistical significance at the 5% and 1% level, respectively. $p - p^*$ is the deviation of housing prices from the cointegrating relation. "LR, $\Delta p_{i,t-1}$" and "LR, $p_i - p_i^*$" refer to the p-value in Likelihood Ratio test on the hypothesis that the coefficients on $\Delta p_{i,t-1}$ and $p_i - p_i^*$, respectively, are the same in every region. The table does not show the intercepts that are estimated for each city separately. The intercepts are available from the authors upon request. Baseline model is the conventional fixed-effects model. In Interaction model 1, all coefficients vary between all cities. Any of the residual series do not exhibit residual autocorrelation based on the Lagrange Multiplier test at lag length one.