Turnover and Price in the Housing Market: Causation, Association or Independence?*

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Abstract

We give general conditions for the data generating process to facilitate the testing of structural dependence of turnover on the average (or median) price in the homogenous housing market. Furthermore, the implications of aggregation over sub-markets is studied. A plausible explanation of the disparate empirical findings in this literature may be aggregation over heterogenous sub-markets. This conclusion is supported by empirical findings using longitudinal quarterly data for 289 Swedish municipalities during 1981:1-2000:2.

Keywords: Aggregation bias, House price, Volume of trade, Search.
JEL Classification: C43, C51, D10

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1 Introduction

Empirical observations of short-run fluctuations in turnover and average prices as well as price-turnover correlations on the second-hand housing market have lately attracted much attention and led to both theoretical and empirical research.[1] These empirical observations question the merits of the widely used model in Poterba (1984).[2] In his model, the turnover level is unaffected during the adjustment process after a shock, since sales occur instantly and without friction. The motive for research is that the presence of frictions in the housing market is associated with costs for the individual as well as for society, since it prevents the price from falling, so that turnover would remain at a constant level. A well-functioning housing market should provide an adequate turnover of residential property of various tenure types, thereby assisting economic agents in accumulating an appropriate asset portfolio to finance life-cycle consumption plans. A housing market with relatively little friction –in terms of shorter sales duration and less costly search – may also assist job mobility and the efficient matching of jobs within the labor market, thereby promoting business enterprises.

Several explanations for the observed fluctuating turnover and the correlation between average prices and turnover have been suggested. Case and Schiller (1989) argued that the rational response in a falling market is to hold on to existing housing investments in anticipation of positive future returns. Stein (1995) suggested that data regularities gave some support for down-payment requirements and other borrowing constraints. Wheaton (1990) suggested that sellers' reservation prices were partly determined by their expectations about sales duration and the cost of holding two units, since sellers were, in fact, buyers seeking to dispose of their old home. Berkovec and Goodman (1996) argued that while the

[1] For instance, using US nation-wide monthly data, covering the period 1968-1992, Stein (1995), found the contemporary price-turnover correlation to be positive, and also that the turnover of houses was positively related to changes in the price level. In periods of a rising (falling) market price, there was also a higher (lower) turnover in the housing markets.

[2] See Åsberg and Åsbrink (1998) and Agell et al. (1998) for a study of the capitalization effects in the market for owner-occupied housing resulting from the 1991 Swedish tax reform. In this framework, house prices are forward looking and solely depend on price fundamentals, i.e., user cost of capital, rents, and construction costs.

contemporary price-turnover correlation should be positive in low frequency data (i.e., annual or quarterly), it might well be negative in high frequency data (i.e., weekly or monthly). If sellers have an information lag as compared to buyers, buyers respond more rapidly to demand shocks. This would cause the number of sales to respond more quickly than the mean prices, due to changes in demand. Therefore, turnover should be superior to the mean price as an indication of changes in demand in high frequency data.

The theoretical research that has tried to explain the observed empirical pattern has focused on the microeconomic underpinning of the observable data. It is therefore important to consider the quality of the aggregated data – turnover and price level – that have been used to test the microeconomic behavior of agents. In empirical tests, the price level is either measured as the average or the median selling price. Hence, the price measure is functionally related to the turnover measure.

In this paper, we (i) discuss some general conditions when the data generating process (DGP) does not result in mean independence between the price level and turnover, due to the functional dependence between these two measures. We also (ii) clarify some of the difficulties in using aggregated data to test the underlying micro behavior of agents in the housing market.

The problem of aggregation has been explored from various viewpoints. Our paper should be viewed as a contribution to the part of the prior literature that examined the effects of failing to empirically incorporate heterogeneity across sub-groups into the aggregate analysis. Some of the prominent authors pursuing this approach are Theil (1954), Grunfield and Griliches (1960) and Stoker (1986). (See Shumway and Davis (2001) for a broad survey of the aggregation literature in general and inferential errors due to aggregation in agricultural economics in particular.)

Our results are the following. (i) The observed price-turnover correlation using aggregated data may be a result of the aggregation. (ii) The sign and the magnitude of the aggregation bias are in general undetermined, since they depend on the cross-sectional heterogeneity in price levels and responses to demand shocks. The average price for houses sold in a specific time period does not only reflect tendencies in the general price level, but will also depend on the composition of houses sold. If, for instance, the number of sales in a high-priced sub-market expands relatively more – as compared to a low-priced sub-market – as a result of an expansion in the demand for housing, then the observed aggregate average price will be positively correlated with the observed aggregate number of sales. (iii) The prospect of finding empirical support for various hypotheses suggested in the literature will depend on the level of aggregation, since aggregation will have a flattening effect on non-zero correlations present at disaggregated levels (i.e. attenuation bias).

The rest of the paper is organized as follows. Section 2 reviews the literature and presents some empirical regularities in the data. In Section 3, the DGP of mean prices and turnover are discussed. The empirical implications and considerations of using aggregated price-turnover data are discussed in section 4. Finally, Section 5 concludes.

2 Review of literature and empirical regularities

Stein (1995) suggested that the positive correlation between house prices and turnover in housing is systematic and could be due to liquidity constraints for the buyer. An initial shock that knocks prices down, weakens the ability of would-be movers to make down payments on new homes. As house prices fall, some potential movers find their liquidity to be so impaired that they are better off staying in their old home than moving to a new, smaller home. Mismatched low-liquidity house-owners may start "fishing"; i.e., listing their house at an above market price and thereby accepting longer waiting times, in a (low-probability) hope of receiving enough money for a reasonable down payment on a new home. Given that the alternative to fishing is not moving at all, the opportunity cost of fishing is low. In times when prices are high, fishing is less attractive since mismatched house-owners can move to the desired home with a shorter waiting time and higher certainty. This should suggest that the trading volume (or the inverse of the length of time on the market) will be positively related to the price level. Not reducing reservation prices sufficiently in falling markets implies that the turnover should fall when the price level is low. Using empirical micro-data, Henley (1998) found that negative housing equity had a negative effect on mobility.

Wheaton (1990) introduced costly search and uncertainty and studied the impact of the cost of holding two units on the behavior of sellers and buyers in a search theoretical framework. He assumed that the number of units and the number of households were fixed in the short run, and that the prospect of remaining mismatched in the current home determined both the search effort and the offer price made by buyers. In his model, sellers are also buyers who have found a new unit and are trying to dispose of their old house. Their reservation prices are determined by expectations about sales time and the cost of holding two units. A higher vacancy rate (i.e., the number of houses up for sale per household) will increase the average sales time (i.e., the inverse of the number of sales). This will also reduce sellers' reservation prices and lead to lower observed market prices. Wheaton also expected a positive relation between prices and the number of sales.

Using monthly US national data from 1968:2-1993:2, Berkovec and Goodman (1996) noted that (i) changes in sales were positively related to changes in aggregated housing demand, that (ii) the positive price-turnover correlation was strongest in low frequency data and that (iii) sales seemed to respond more quickly to changes in housing demand than the price level. Berkovec and Goodman argued that these observations were consistent with sellers having an information lag as compared to buyers and that buyers therefore responded more rapidly to demand shocks. Initially, a demand shock would only affect the number of buyers or their willingness to pay. Hence, turnover would react more quickly to changes in demand than aggregate price movements. The predictive power of turnover in measuring demand fluctuations would therefore be superior to that of prices, at least in high frequency data, where there might well be a negative price-turnover correlation.[3]

The empirical results reported in the literature are not uniform, however. Follain and Velz (1995) reported that house prices and turnover were negatively related. Their empir-

[3] Hort (2000) found support for the statements made by Berkovec and Goodman in Swedish data. To test the hypothesis of a slow adjustment for sellers, she estimated impuls-response functions from a three-dimensional VAR model, with the after-tax mortgage rate, turnover and price. In monthly, but not in quarterly, data turnover reacts faster than prices to a shock in the after-tax mortgage rate. She did not perform her analysis on data of lower frequency, however.

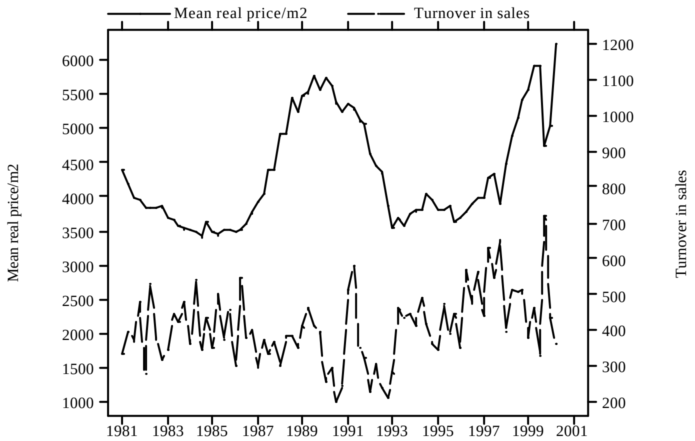

Figure 1: Mean real selling price per square meter (nominal price/CPI 1980) and turnover in sales of single-family houses, Stockholm municipality, 1981:Q1-2000:Q2

ical approach differs from Stein's in two ways. They estimated a structural model of the housing market with four simultaneously determined equations (real constant-quality price level, turnover, housing quality index, and housing stock). Furthermore, they used annual disaggregated data, spanning over 22 US metropolitan areas during 1986-1992.

2.1 Some empirical evidence for Sweden

For the two major metropolitan municipalities in Sweden, we can also find some support for a negative correlation between the level of housing prices and the level of sales. Figures 1 and 2 show the mean real price per square meter ($\overline{p}_t$) and turnover (q_t) in the single-family housing market in Stockholm and Gothenburg, respectively.[4] The raw correlation coefficients for quarterly data (1981:1-2000:2) are -0.12 for Stockholm and -0.25 for Gothenburg. The correlations are most pronounced in the time period before 1991 with -0.37 for Stockholm and -0.51 for Gothenburg while, in the period after 1991, the correlations are almost zero. The raw correlations between the recent change in house prices and turnover level are,

[4] The data used in this chapter was collected by the National Land Survey of Sweden (Lantmäteriet). The data cover almost one million sales of single-family houses. It consist of the mean per square meter selling price and the number of sales for single-family houses in all of 289 municipalities. The data frequency is quarterly and the time period is 1981:1-2000:2.

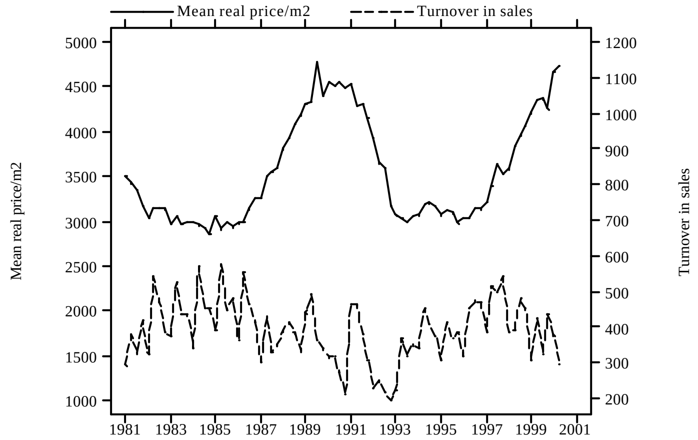

Figure 2: Mean real selling price per square meter (nominal price/CPI 1980) and turnover in sales of single-family houses, Gothenburg municipality, 1981:1-2000:2

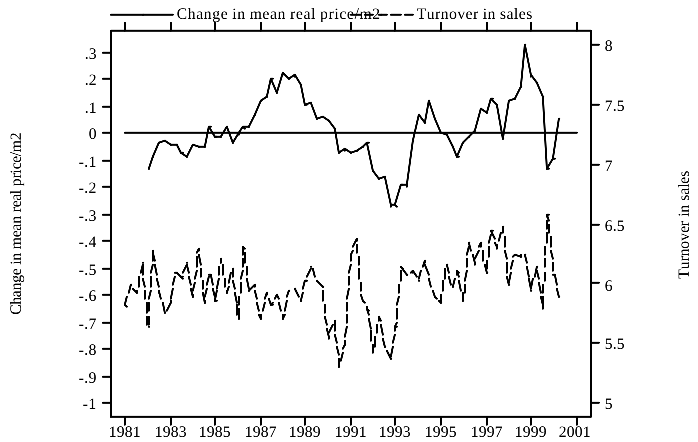

Figure 3: Percentage annual change in mean real selling price per square meter (nominal price/CPI 1980) and (the log of) turnover in sales of single-family houses, Stockholm municipality, 1981:1-2000:2

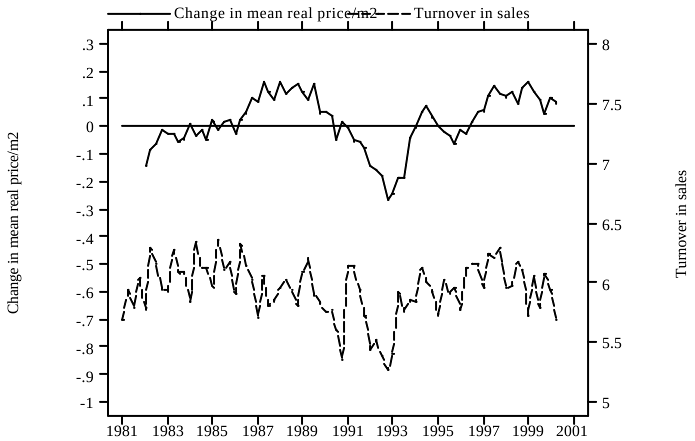

Figure 4: Percentage annual change in mean real selling price per square meter (nominal price/CPI 1980) and (the log of) turnover in sales of single-family houses, Gothenburg municipality, 1981:1-2000:2

however, positive, with 0.21 for Stockholm and 0.34 for Gothenburg. These latter correlations support Stein's (1995) hypothesis, but that is not the case with the first correlations based on the levels. Figures 3 and 4 plot the annual percentage change in real house prices per square meter and the number of sales (in logs) for the two municipalities.[5]

The raw correlation is, however, not without demur when it comes to drawing convincing inference about, e.g., Stein's (1995) hypothesis. First, the raw correlation does not control for a number of background variables. Second, one might suspect that there is a substantial amount of heterogeneity across regions and time periods for which the raw correlation above does not account.

In order to, at least partly, accommodate these objections, we run a series of regression models with various specifications. First, we regress the mean real price per square meter (in logs) on turnover (also in logs) (henceforth level-vs.-level). In our second model specification, we instead use the annual percentage change in real price per square meter as the dependent variable (henceforth difference-vs.-level). These specifications explore data along the lines of Stein's (1995) argument. In the third model specification, we regress the annual percentage

[5] The annual percentage change of the mean real price is computed as $\ln \overline{p}_t - \ln \overline{p}_{t-4}$.

Table 1: Slope estimates from models with one common slope parameter using municipal (m) and county (c) data. Time period 1981:1-2000:2.

	Municipality data			County data
	(1)	(2)	(3)	(4)
	$\ln \bar{p}_{mt}$	$\ln \frac{\bar{p}_{mt}}{\bar{p}_{mt-4}}$	$\ln \frac{\bar{p}_{mt}}{\bar{p}_{mt-4}}$	$\ln \bar{p}_{ct}$
$\ln q_{mt}$	0.017 (3.04)			
$\ln q_{mt}$		0.016 (2.60)		
$\ln \frac{q_{mt}}{q_{mt-4}}$			0.016 (3.20)	
$\ln q_{ct}$				0.014 (0.94)
Observations	22206	21015	21015	1632
N		289		21

Note: Robust t-values in parentheses. Quarterly effects and fixed time effects included. (1)-(3): The number of parameters (including a municipal factor) are 367. (4): The number of parameters (including a county factor) are 99.

change in real price per square meter as the dependent variable and the annual percentage change in turnover as the independent variable (henceforth difference-vs.-difference).[6]

The estimated slope coefficients (and t-values) are given in Table 1. We find the slope coefficient estimates to be positive and significant, and almost exactly the same in all three model specifications. According to these results, data seem to give a very uniform picture of the price-turnover correlation. The positive estimate in columns (1) and (2) seem to confirm Stein's (1995) finding of a positive price-turnover correlation and that the turnover of houses and recent changes in the price level are positively related.

However, a quite different and very irregular description of the data concerning these relationships arises if we drop the restriction of a single common slope parameter for all municipalities, and instead allow each municipality to have its own separate slope coefficient.[7]

[6] In all specifications, mean price and turnover are measured for each municipality. In the regressions, we include all 289 municipalities for 1981:1-2000:2. Furthermore, we include quarterly fixed effects to adjust for seasonal effects, fixed municipality intercepts to allow for cross-sectional heterogeneity in the price levels, and fixed time effects to control for the influence of common demand and supply conditions.

Figure 5 plots the estimated municipality-specific slope coefficients from the level-vs.-level specification against the county affiliation code.[8] Here the significance level is indicated at the 5 percent level. It is clear that the model with a uniform slope parameter is a feeble description of data. For the greater part of the sample, the turnover level has no significant influence on the price level. Only 22 out of the 289 slope estimates are significantly positive and 35 significantly negative at the 5 percent level. Furthermore, the spread in the slope estimates varies substantially over different parts of the country. The spread of the estimates for the municipalities mainly located in the western part of the country (county affiliation codes 14, 17 and 18) is considerable. The significant estimates are both positive and negative, but there is a clear dominance towards positive ones. In the rest of Sweden, significant estimates are mainly negative. In the southern parts of the country (county affiliation codes 10 and 12), all significant estimates are negative and some are quite large.

We repeat this exercise for the difference-vs.-level and difference-vs.-difference specifications (see Figures 6 and 7, respectively, with the significance level is indicated at the 5 percent level). Once more, we conclude that there is a substantial amount of dispersion in these relationships, and that the uniform estimates given in Table 1 give a fairly poor description of these relationships in the data. Almost all slope estimates are insignificant. In the difference-vs.-level regression, we note that almost all significant estimates represent a positive relationship (33 significantly positive and 3 significantly negative estimates). In the difference-vs.-difference specification, we find a slight dominance towards positive ones (28 significantly positive and 20 significantly negative estimates). The dispersion in the estimates over different parts of the country from both these specifications is similar to that

[7] The model is a least square dummy variable model. In this regression, we interact turnover with the municipality dummies, but otherwise keep the specification intact as compared to that in Table 1. A robust covariance estimator (White, 1980) is used to calculate the t-values. The number of observations is 22206 and the number of parameters is 655.

[8] The county affiliation code classification is Stockholms län (1), Uppsala län (3), Södermanlands län (4), Östergötlands län (5), Jönköpings län (6), Kronobergs län (7), Kalmar län (8), Gotlands län (9), Blekinge län (10), Skåne län (12), Hallands län (13), Västra Götalands län (14), Värmlands län (17), Örebro län (18), Västmanlands län (19), Dalarnas län (20), Gävleborgs län (21), Västernorrlands län (22), Jämtlands län (23), Västerbottens län (24), Norrbottens län (25).

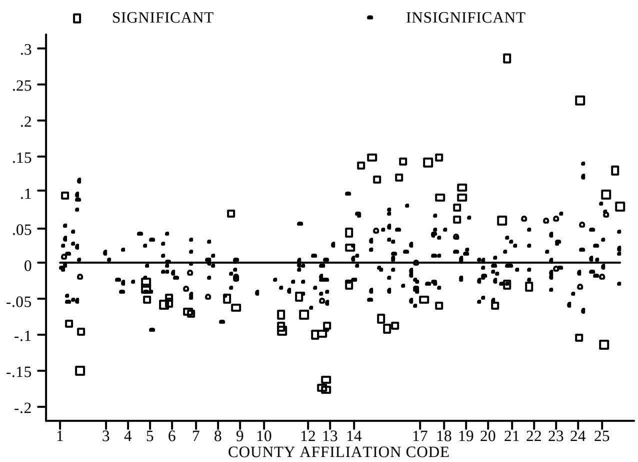

Figure 5: Slope coefficients, significant (□) on the 5 percent level and insignificant (○), on the y-axis, from a regression with real price per square metre (in logs) as the dependent and turnover (in logs) as the independent variable, and the county code on the x-axis (level-vs.-level specification).

of the level-vs.-level specification.

Taken together, these exercises clearly show that there is a substantial amount of heterogeneity between different parts of Sweden, and that neglecting these irregularities may cause incorrect inferences. Considering the spread and the small part of significant estimates in Figures 5, 6 and 7, it is likely that the results in Table 1 are driven by a relatively small fraction of outliers, and that they depend on how we divide the sample into different time periods and which regions of the country are included.

Different authors have used different levels of aggregation to study these relationships (cf. the review in this section). However, the level of aggregation will influence inference. We illustrate this by aggregating the municipal data, which we used above, up to the county level. In the latest county classification, Sweden's 289 municipalities are geographically classified into 21 counties. We re-estimate our level-vs.-level model with the same model structure as before that allows for county-specific slope parameters, and instead use the county level data.[9]

[9] The aggregated mean price in county c at time t is calculated as $\bar{p}_{ct} = \sum_m q_{mt}\bar{p}_{mt} \; / \sum_m q_{mt}$, and the

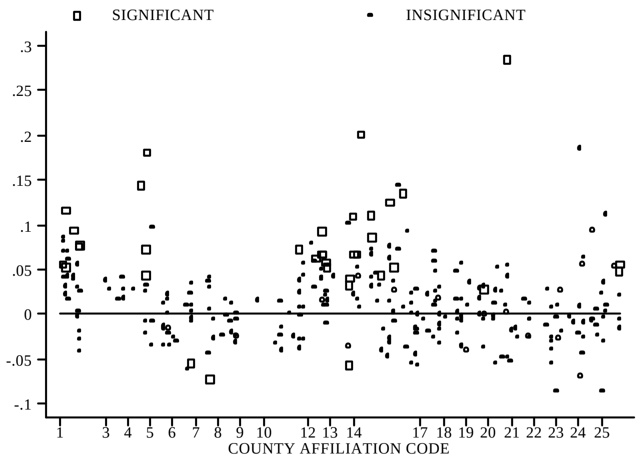

Figure 6: Slope coefficients, significant on the 5 percent level (□) and insignificant (○), on the y-axis, from a regression with the annual percentage change in real price per square metre as the dependent and turnover (in logs) as the independent variable, and the county code on the x-axis (difference-vs.-level specification).

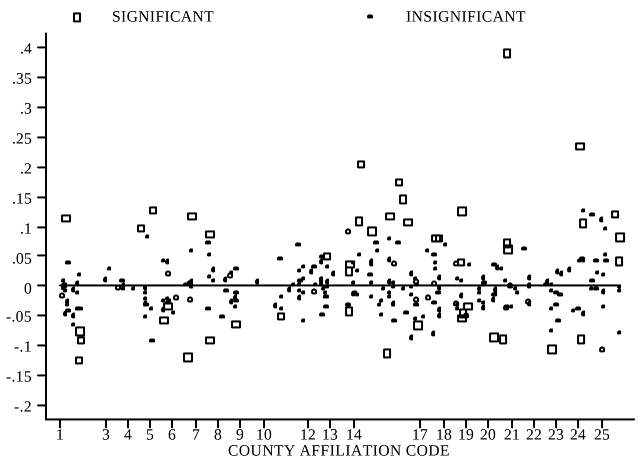

Figure 7: Slope coefficients, significant on the 5 percent level (□) and insignificant (○), on the y-axis, from a regression with the annual percentage change in real price per square metre as the dependent and the annual percentage change in turnover as the independent variable, and the county code on the x-axis (difference-vs.-difference specification).

We present the estimated slope coefficients from this exercise in Figure 8. If inference is more or less independent of the aggregation level, the relative position of estimates and levels of significance from this regression should be roughly the same as those presented in Figure 5. A few things can be noted. First, it is obvious that in absolute terms, there is much less dispersion at the county level than at the municipal level. The aggregation over municipalities smoothens the estimates found at the municipality level, which is what should be expected due to the influence of the attenuation bias. Second, a larger share of the slope coefficient estimates in the county sample are significant as compared to those in the municipality sample. For 21 counties, 6 estimates are significantly negative, 1 is significantly positive and the rest (14 counties) are insignificant. Third, the varying dispersion of significant estimates over different parts of the country only agrees with what was found at the municipal level in some cases (in Figure 5). The negative (and significant) estimates in the southern parts of the country (county affiliation codes 10 and 12) match relatively well with what was found when using the municipality data. However, the estimates for the western parts of the country (county affiliation codes 14, 17 and 18) are practically zero. When using the municipality level, we found quite high and positive estimates in those regions which are obviously wiped out in the more aggregated data. Fourth, estimating a common slope for all counties using the county level results in a somewhat lower estimate as compared to the corresponding estimate on municipality level data. The point estimate is positive, however, insignificant (See Table 1, column (4), for details).[10]

aggregated turnover is $q_{ct} = \sum_m q_{mt}$, where $\bar{p}_{mt}$ is the mean price in municipality m at time t, and q_{mt} is the turnover in municipality m at time t. We estimate the level-vs.-level model on a county level with the real price per square meter (in logs) as the dependent variable and turnover (also in logs) interacted with the county affiliation code as independent variables. We include dummies for quarter, county and time. A robust covariance estimator (White, 1980) is used to calculate t-values. The number of observations is 1632, and the number of parameters is 119.

[10] The result in Hort (2000), who also used Swedish county level data, deviates from our result in this respect. She presented (in Table 1) a negative and significant point estimate from a model with one common slope parameter for all counties. She used quarterly data from a slightly different time period than ours, however (1982:3-1996:2). She used fixed time and fixed county effects, but did not include fixed seasonal effects. Instead, she used data series on turnover, which were seasonally adjusted series by cross-sectional unit-specific seasonal dummies. Constant quality house price indexes were also used (see Hort (2000) for details).

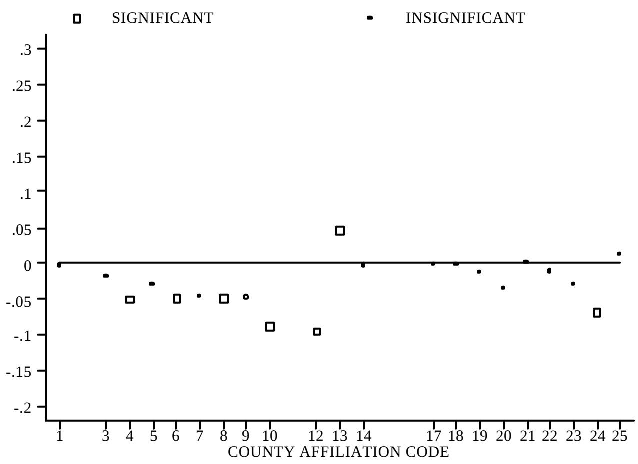

Figure 8: *County data.* Slope coefficients, significant on the 5 percent level (□) and insignificant (○), on the y-axis, from a regression with the real price per square metre (in logs) as the dependent and turnover (in logs) as the independent variable, and the county code on the x-axis (level-vs.-level specification).

Once more, the time-period chosen for study and the counties included would affect the results. It is, however, obvious that aggregation can substantially change inference. Our example of aggregating the price and turnover data from the municipality level to the county level shows the familiar result of an attenuation bias in aggregated data. Next, we discuss the DGP for the average price and turnover. In the following section, we give analytical results of the aggregation bias of the ordinary least squares (OLS) estimator, given the assumptions of the DGP.

3 Price and quantity formation

To validate different hypotheses about the microeconomic behavior of agents in the housing market (e.g., asymmetric demand shock responses of buyers vis-à-vis sellers (Berkovec and Goodman, 1996), or costly search and uncertainty (Wheaton, 1990)), it is most common for researchers to test $\beta \neq 0$ in one of the three model specifications given by

$$\vec{p}_t^* = \alpha(\mathbf{x}_t) + \beta q_t^* + \varepsilon_t, \tag{1}$$

where $\{\bar{p}_t^*, q_t^*\}$ is either $\{\bar{p}_t, q_t\}$, $\{\Delta\bar{p}_t, q_t\}$ or $\{\Delta\bar{p}_t, \Delta q_t\}$, and Δ is the time difference operator. Here, q_t is the turnover and $\bar{p}_t = q_t^{-1}\sum_{j=1}^{q_t} p_{jt}^R$ is the average price, where p_{jt}^R is the realized selling price of the j:th house sold during a specific time period. x_t is a vector of explanatory variables, which are presumably correlated with q_t. For simplicity, we let $\alpha(x_t) = \alpha$ in the following. With reference to the former section, these three specifications hence correspond to the level-vs.-level, difference-vs.-level and difference-vs.-difference specifications except that $\bar{p}_t$ and q_t are not in logarithms. Note that α and β are not necessarily the same in the three specifications.

Consider estimating (1) using OLS. The mean independence of q_t or Δq_t and ε_t is sufficient for the OLS estimator to be consistent (i.e. $E(\varepsilon_t|q_t) = E(\varepsilon_t|\Delta q_t) = 0$). Since the mean price is functionally dependent on the number of houses sold, it is of interest to discuss the assumptions needed on the DGP to implement tests of structural dependence using the above specifications, i.e. what assumptions on the DGP are needed for q_t or Δq_t to be mean independent with ε_t in (1) above?

The frequent assumption in the literature about the DGP is that a house sale is not frictionless. A search theoretical framework is quite often used to model the process of a house sale; it is costly and takes time to carry out since it takes time for a seller to find the right buyer, for a buyer to find the right house, and for the seller and buyer to agree upon the price.

Consider a housing market with a constant and homogeneous housing stock H, where each house-owner owns one unit of the housing good. At each t, a house-owner decides whether to enter the market and become a seller at a given separation rate. Seller j sets a reservation price, π_{jt}. The distribution of reservation prices at t is described by $F_\Pi(\pi_t)$.[11] Assume that the decision to sell cannot be revoked.

The bid distribution (generated by the population of buyers) is $F_P(p_t)$. Bids are assumed

[11] The reservation price π_{jt} can be derived - under assumptions of exogenous separation and offer rate - using a search theoretical framework. The spread in the reservation price distribution $F_\Pi(\pi_t)$ can be motivated by the fact that sellers may be heterogeneous in their motivation to sell (see also Glower et al., 1998), or by duration dependence (see, e.g., Sirmans et al., 1995).

to arrive at the offer rate λ_t. A seller sells if the incoming bid is greater than his reservation price. The resulting price realization when a sale goes through, p_{jt}^R, depends on the buyer's and seller's information about the market. We will discuss this below.

To enhance the presentation, let $\beta \equiv 0$. Now, if $E(\bar{p}_t|q_t) = E(\bar{p}_t)$ (or $E(\bar{p}_t|\Delta q_t) = E(\bar{p}_t)$), then ε_t and q_t (or Δq_t) in (1) are mean independent. Thus, the bias of the OLS estimator is zero. The conditional expectation of $\bar{p}_t$ is

$$E(\bar{p}_t|q_t) = \frac{1}{q_t}\sum_{j=1}^{q_t} E(p_{jt}^R|q_t).$$

Consequently, if $E(p_{jt}^R) = E(p_t^R)$ for all $j = 1, ..., q_t$, then $E(\bar{p}_t|q_t) = E(\bar{p}_t)$ and the bias of the OLS estimator is zero. The meaning of the condition $E(\bar{p}_t|q_t) = E(\bar{p}_t)$ is that the ordering of the sales is of no importance. Selling an additional unit or a unit less, given that some units are sold, does not alter the expected mean selling price. If the ordering is of importance for $E(\bar{p}_t|q_t)$, then $\bar{p}_t$ and q_t are mean dependent, even if $\beta \equiv 0$.

We can express this formally by ordering the sales in the natural order, with $j = 1$ being the first house sold and $j = q_t$ the last house sold during a specific time period. The unconditional expectation of $\bar{p}_t$ is thus

$$\begin{aligned} E(\bar{p}_t) &= \sum_{k=1}^{\infty} \Pr(k) E(\bar{p}_t|k) = \sum_{k=1}^{\infty} \Pr(k) \frac{1}{k} \sum_{j=1}^{k} E(p_{jt}^R|k) \qquad (2) \\ &= \sum_{k=1}^{\infty} \frac{1}{k} \left(\Pr(k) E(p_{kt}^R|k) + \Pr(k|k-1) \sum_{j=1}^{k-1} E(p_{jt}^R|k-1) \Pr(k-1) \right), \end{aligned}$$

where $E(p_{jt}^R|k)$ is the expected value of the j:th house sold, given that k sales occurred, and $\Pr(k)$ is the probability that k sales occur. Thus, if $E(p_{jt}^R|k)$ is independent of the ordering of sales, the expected value of the last house sold is not different from that of the houses sold before. The average price and turnover are mean independent and $E(\bar{p}_t|q_t) = E(\bar{p}_t)$. Whether this is true in reality is an open question. It will depend on the friction in the homogenous market; how well-informed buyers are about the reservation prices they meet and how well-informed sellers are about incoming bids.

Consider the extreme situation where a buyer has perfect knowledge about individual reservation prices. He would place his bid $p_{it} = \pi_{jt}$ and buy immediately. Then, the matching is not random, since buyers would search for the cheapest possible house. The first house to be sold would be the cheapest, the second house would be the second cheapest, and so on. So for each k, we would get $E(p^R_{1t}|k) < E(p^R_{2t}|k) <, ..., < E(p^R_{kt}|k)$,[12] and hence $E(\bar{p}_t|q_t) \neq E(\bar{p}_t)$ and the bias of the OLS estimator is different from zero.[13] Observe that this extreme situation of asymmetric information and no market is not likely to occur in real life. We have assumed that there is no competition for houses on the market and that buyers are perfectly informed about all prices, whereas sellers are not.

A more relevant situation is if sellers announce list prices, or asking prices, which are known to the buyer. (A list price indicates the seller's intent, and gives the buyer further, but not perfect, knowledge about π_{jt}.)[14] This situation is likely to introduce mean dependence between ε_t and q_t or Δq_t, hence biasing the estimate of β above.

Thus, even though the average price and turnover are functionally related, it is possible to test structural relations, given specific restrictions on the DGP. If every house-seller meets the same distribution of bids and the offer rate λ_t is the same for all sellers, the bid size and the bid offer rate are independent. If buyers pay according to their personal valuation of the house (i.e. auctions) then, for a randomly chosen sale, $E(p^R_{jt}) = \int E(p_{it}|p_{it} > \pi_t)dF_\pi(\pi_t)$, i.e. the mean accepted offer price. If buyers have perfect knowledge of individual reservation prices, the corresponding expected price is the mean reservation price, i.e. $E(p^R_{jt}) = E(\pi_t)$. It is worth noting that if the reservation price distribution is degenerated (i.e. one point) and $E(p^R_{jt}) = E(\pi_t)$ (either because buyers do not pay according to their personal valuation or the reservation price is known), then ε_t and q_t or Δq_t are mean independent.

[12] Holds if F_Π is not a single-point distribution.

[13] Introducing list prices into the model would alter the seller's bid distribution and the probability of receiving an offer, as argued in Horowitz (1992). This would introduce a correlation between the offer rate and bids and create a dependence between turnover and the mean selling price.

[14] Horowitz (1992) argued that list prices appeared to be price ceilings above which no bids arrived, and that higher list prices reduced the probability of recieving an offer.

In this section, we have – given a quite general framework of price and turnover formation – discussed the possibilities of testing structural hypotheses based on β in the regression models above and the implications for the DGP. One implication is random shopping which was, for instance, the identifying assumption in Berkovec and Goodman (1996) when they tested a hypothesis of slow adjustment of sellers vis-à-vis buyers. In their model, the reservation price distribution was obtained from a negative market duration dependence of the reservation price. Under the assumption of random shopping, other structural hypotheses can be tested using the above regressions.[15] ,[16] Furthermore, note that if we exchange the mean price for the median price (p_t^m) in equation (2), we will obtain the same results as above.[17]

4 Empirical considerations

In this section, we will consider estimations and testing of the structural relationships (1) using aggregated data. Assume that we have data on prices and turnover at the municipal level and that the housing market in the municipality consists of A different, but perfectly homogenous, areas (a). Let p_{kat}^R be the realized price in sale $k = 1, ..., q_{at}$, in area a in the time period $t = 1, ..., T$. Then, the average price at the aggregate municipal level is

$$\overline{p}_t = \frac{\sum_{a=1}^{A} \sum_{k=1}^{q_{at}} p_{kat}^R}{q_t} = \frac{\sum_{a=1}^{A} q_{at} \overline{p}_{at}}{q_t} = \sum_{a=1}^{A} \phi_{at} \overline{p}_{at}, \tag{3}$$

[15] The presence of non-random shopping (i.e. some form of asymmetric information or no market) is naturally a structural hypothesis that it would be of interest to test.

[16] Observe that ε_t and q_t are statistically dependent also under the assumption of random shopping, although mean independent. This is easily seen from calculating the variance: $Var(\varepsilon_t) = \sigma^2 q_t^{-1}$, where $\sigma^2 = Var(\varepsilon_{jt})$.

[17] This follows by exchanging $\bar{p}_t$ with p_t^m in equation 2 above, hence

$$\begin{aligned} E(p_t^m) &= \sum_{k=1}^{\infty} \Pr(k) E(p_t^m | k) = \sum_{k=1}^{\infty} \Pr(k) E(p_{kt}^{Rm} | k) \\ &= \sum_{k=1}^{\infty} \Pr(k) E(p_{kt}^{Rm} | k) + \Pr(k | k-1) \sum_{j=1}^{k-1} E(p_{jt}^{Rm} | k-1) \Pr(k-1), \end{aligned}$$

where p_{kt}^{Rm} is the observed median price, given k sales.

where $\overline{p}_{at} = q_{at}^{-1} \sum_{k=1}^{q_t} p_{kat}^R$ is the average price in area a, p_{kat}^R is the realized selling price of the k:th house sold in area a (where $E(p_{kat}^R) = \mu_a$, according to Section 3, and μ_a is a parameter), $q_t = \sum_{a=1}^{A} q_{at}$ is the aggregate turnover, and $\phi_{at} = q_{at}/q_t$ is the proportion of total sales in area a.

Let us consider the problem of aggregate data by specifying a simple structural relationship at the area level. The average price level in area a is specified as

$$\bar{p}_{at} = \alpha_a + \beta q_{at} + \varepsilon_{at},$$

where q_{at} is the turnover.[18] Observe that if $\beta = 0$, then $\alpha_a = \mu_a$. Here ε_{at} is, by assumption, mean independent of q_{at}. Hence, if data are available at the disaggregated level, an unbiased inference using OLS is possible.

Now, consider estimating

$$\overline{p}_t = \alpha + \beta q_t + \varepsilon_t$$

with OLS. Let $q_. = \frac{1}{T} \sum_{t=1}^{T} q_t$ be the mean turnover for the period studied and assume that $Q_{qq} = \text{plim} \frac{1}{T} \sum_{t=1}^{T} (q_t - q_.)^2$ exists. The OLS slope estimate will converge in probability to

$$\begin{aligned} \text{plim}\, \widehat{\beta} &= (Q_{qq})^{-1} \,\text{plim}\, \frac{1}{T} \sum_{t=1}^{T} (q_t - q_.) \overline{p}_t \\ &= (Q_{qq})^{-1} \,\text{plim}\, \frac{1}{T} \sum_{t=1}^{T} (1 - \frac{q_.}{q_t}) \left[\sum_{a=1}^{A} \alpha_a q_{at} + \beta \sum_{a=1}^{A} q_{at}^2 \right] \\ &= \delta + \kappa \beta. \end{aligned} \tag{4}$$

Here

$$\delta = (Q_{qq})^{-1} \,\text{plim} \frac{1}{T} \sum_{t=1}^{T} (q_t - q_.) \sum_{a=1}^{A} \phi_{at} \alpha_a, \tag{5}$$

and

$$\kappa = (Q_{qq})^{-1} \,\text{plim}\, \frac{1}{T} \sum_{t=1}^{T} (1 - \frac{q_.}{q_t}) \sum_{a=1}^{A} q_{at}^2. \tag{6}$$

[18] The result from aggregation when using this model can easily be extended to the other two models (given in equation 1). The result from this is qualitatively the same as the results below.

Since $\sum_{t=1}^{T}(q_t - q_.)^2 = \sum_{t=1}^{T}(1 - q_./q_t)q_t^2$ and

$$(1 - \frac{q_.}{q_t}) \sum_{a=1}^{A} q_{at}^2 < (1 - \frac{q_.}{q_t}) q_t^2, \tag{7}$$

it is evident that $\kappa < 1$. Hence, if $\beta \neq 0$, the OLS estimate of the slope will be attenuated by aggregation.[19] It is interesting to note that the aggregation bias occurs for two different reasons: *attenuation bias* emerges since $\kappa < 1$ and *structural bias* (i.e. $\delta \neq 0$) arises if the areas are heterogeneous in (i) price levels (i.e. $\alpha_a \neq \alpha$ for some $a = 1, ..., A$) and (ii) responses to market shocks. Thus, even if there is no structural dependence (i.e. $\beta = 0$), the OLS estimator may be biased.

Let us exemplify with two areas ($A = 2$). Then, since $\phi_{1t} = 1 - \phi_{2t}$, we can write

$$\begin{aligned} \delta &= (Q_{qq})^{-1} \operatorname{plim} \frac{1}{T}(\alpha_2 - \alpha_1) \sum_{t=1}^{T} (q_t - q_.)\phi_{2t} \\ &= (\alpha_2 - \alpha_1) C_{q\phi} V_{qq}^{-1} = (\alpha_2 - \alpha_1) \widehat{\beta}_{q\phi}, \end{aligned}$$

where $C_{q\phi}$ is the covariance between ϕ_{2t} and q_t, V_{qq} is the variance of q_t and $\widehat{\beta}_{q\phi}$ is the OLS parameter in a regression of ϕ_{2t} on q_t.

Hence, if the expected price levels in the two areas are equal $(\alpha_1 = \alpha_2)$, or the proportions of the turnovers for the areas (of the total turnover) remain unchanged over the period studied, there will be no structural bias.

Suppose instead that area 2 consists of high-priced housing and area 1 of low-priced housing, i.e. $\alpha_2 > \alpha_1$, and that the overall turnover increases gradually over time due to a positive demand shock[20] and that this increase is mainly due to a gradual increase in the sales in the high-priced area 2, i.e. $\widehat{\beta}_{q\phi} > 0$. Then, the estimate will be positively biased i.e., $\delta > 0$. The intuition is that the high-priced area 2 gradually attains a higher weight in the aggregated mean price function. Consequently, q_t and $\overline{p}_t$ will be positively correlated, even though q_{at} and $\overline{p}_{at}$ are mean independent. If, instead, trade in the high-priced area 2

[19] Note that the estimated correlation may even change signs since κ is not bounded from below.

[20] This shock can e.g. be caused by changes in user cost and exogenous changes in demand (e.g. due to changes in the demographic factors).

increases at a slower rate as compared to the low-priced area 1, then, by the same reasoning, q_t and $\overline{p}_t$ will be negatively correlated and $\delta < 0$. The structural bias can be exemplified with a similar reasoning for $A > 2$ and for the other two model specifications.[21]

Considering equation (4), it can be noticed that the results of aggregation bias do not depend on how the aggregate price is calculated (i.e. using the mean or the median price). This is the case since δ results from ϕ_{at} and q_t being correlated (and $\alpha_a \neq \alpha$ for some $a = 1, ..., A$) and κ stems from the inequality (7). Hence, both κ and δ are independent of the aggregate price measure.

Figure 9 gives an illustration of the aggregation bias for, arbitrarily set, $\delta = 0.2$ and $\kappa = 0.6$. In the origin of the coordinates, there is no structural dependence at the disaggregated homogeneous level. The point estimate is, however, positive since market shocks have a relatively larger effect on high-priced than on low-priced areas. Furthermore, we can see that the size of the bias is a function of the true slope parameter. The bias is reduced (increased) if the true structural dependence between the average price and turnover at the disaggregated level is positive (negative). If the true slope is 0.5, these parameters suggest that there is no bias. However, had we chosen δ to be -0.2, thereby assuming that market shocks have a relatively lower effect on high-priced as compared to low-priced areas, the bias would be zero for $\beta = -0.5$. The regression estimates based on aggregated data may hence not be very informative about the true structural dependence.

It can furthermore be observed that

$$\varepsilon_t = \sum_{a=1}^{A} \phi_{at}\varepsilon_{at} = q_t^{-1} \sum_{a=1}^{A} \sum_{j=1}^{q_{at}} \varepsilon_{jat},$$

where, by assumption, ε_{jat} is identical independent distributed (iid) with the variance σ_a^2 and hence,

$$Var(\varepsilon_t) = q_t^{-1} \sum_{a=1}^{A} \sigma_a^2.$$

[21] The derivation of the bias above naturally depends on the DGP at the area level. Hence, the results from aggregation using the two other specifications depend on the assumed DGP. If the DGP has the same functional form as the model estimated at the aggregate level, then exactly the same result is obtained on the bias as above.

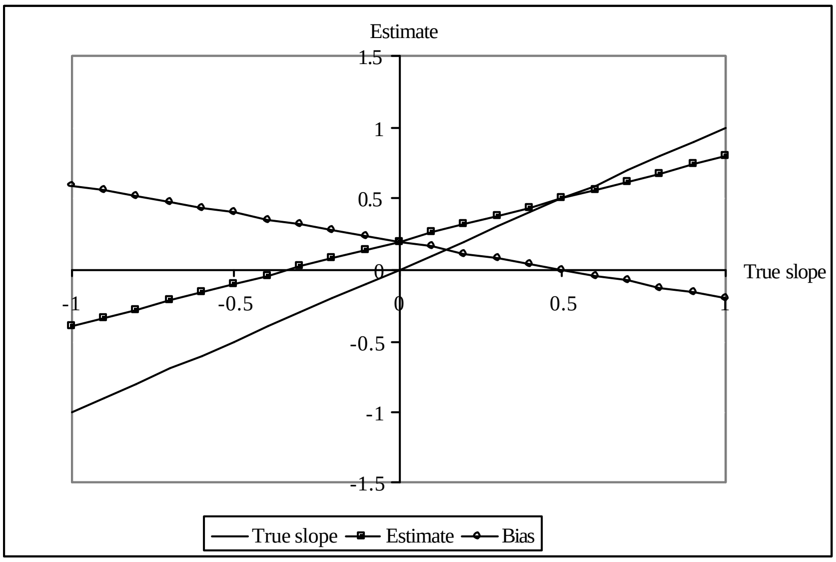

Figure 9: The estimate $\hat{\beta}$, the true β and the bias as a function of β with $\delta = 0.2$ and $\kappa = 0.6$.

Thus, we will have heteroscedasticity and a feasible GLS estimator or a robust covariance estimator should be used in order to obtain the correct sizes.

5 Concluding remarks

This paper discusses the properties of macro economic data – the price level and the number of sales – in the housing market and the consequences of aggregation. In the literature, structural hypothesis tests of the contemporary relationship between the price level and turnover (or between functions of these measures) are most often implemented using linear regression analysis. However, considering that the price level is functionally dependent on turnover since it is either measured as the average or the median selling price, it important to consider and discuss the realism of the implications necessary for the DGP for implementing tests of structural dependence.

It is possible to test structural hypotheses using the contemporary relationship between the price level and turnover, given a quite general search theoretical framework where buyers shop around among houses up for sale in a random fashion (i.e. all house sellers are assumed to receive bids – from the same distribution of bids – at random). Data on prices and turnover may, however, only be available for a large market, where the assumption of random shopping can then be questioned. The regression estimator using these data for estimating the price turnover dependence is consequently likely to be biased. Another way of studying the consequences of non-random shopping is to consider it as an aggregation problem. We can assume random shopping to hold at some disaggregated level.[22] The data may, however, only be available at the aggregate level, which will lead to aggregation bias. The two problems – realism of the assumptions necessary for the DGP for implementing tests of structural dependence and aggregation – are really two different ways of studying the same issue: when is it possible to use aggregated data to test for structural dependence?

The correlation between aggregated data on the price level and turnover has been empirically investigated by several authors at different levels of aggregation, both over cross-sectional units and time periods, but the reported results are not very uniform (cf. Section 2). In this paper, we suggest that a plausible explanation of the disparate empirical findings in the literature is the low quality of data caused by aggregation over heterogenous areas. In line with the prior aggregation literature, we show that aggregation bias is induced if the aggregate analysis fails to empirically incorporate heterogeneity across sub-groups. The theoretical suggestions are supported by empirical findings using Swedish longitudinal data of 289 municipalities during 1981:1-2000:2, containing the mean real price per square meter and turnover in the single-family housing market. There is a substantial amount of heterogeneity between different parts of Sweden and over time, and neglecting these irregularities may cause the wrong inference.

The motive for research has been to gather knowledge of the functionality of the housing market since frictions are associated with high costs for the individual as well as for society in terms of sales duration and costly search. The observed correlations (both negative and positive) between the price level and sales found in the literature may hence not necessarily

[22] In order to make the assumption of random shopping reliable, the size of disaggregated areas can be made as small as necessary. Each house may in fact constitute an area.

be due to different forms of inefficiency in the second-hand housing market as has been suggested by the authors. It may equally well be the effect of an aggregation of heterogenous sub-markets.

The average price for houses sold during a specific time period does not only reflect tendencies in the general price level, but also depends on the composition of houses sold. Data for the homogeneous area would enable the researcher to test structural relationships without restrictive assumptions about the DGP. It is, however, reasonable to assume that the turnover is low, since the size of such areas is likely to be small, thereby causing estimates of the relationship to be imprecise. One way of gaining efficiency would be by pooling the data over the areas and making use of estimated hedonic prices (see, e.g. Englund, et al., 1998) to accommodate for heterogeneity (between the areas) in the price levels.

However, both the assumption of homogenous areas in empirical data and the methods of hedonic price functions can be questioned; the former since homogeneous areas (of a reasonable size) simply do not exist and the latter for not accommodating all heterogeneity. There are also other concerns – connected to the aggregation problem, however – with the use of aggregated data. Spatially distinct housing markets arise from geographical fixity as well as from heterogeneity in physical characteristics of houses and in local ways of providing public goods (see Fratantoni and Schuh, 1999). The price level and turnover dependence can, therefore, be caused by a vast number of structural relations that may also change over time or across pooled units. The reliability of macro economic data can be further questioned, considering the empirical evidence suggested in several studies of influential patterns of contiguity dependence.[23]

There are different forms of inefficiencies in the second-hand housing market. It may be difficult to study one kind of inefficiency using aggregated data simply because other

[23] See Burgess and Profit (2001) for a study of spatial externalities in the labor market. They studied externalities between neighboring travel-to-work areas with respect to the matching of workers and firms in Great Britain and found significant spill-over effects between such areas. Pollakowski and Ray (1997) studied price-ripple effects in the US housing market. Muellbauer and Murphy (1997) and Meen (1997) studied this issue for the US. Berg (2000) studied Sweden, and found significant price ripple effects between regional housing markets.

inefficiencies may attenuate the effect of, e.g., changes in user costs. Take the tax system as an example. In Sweden, there is a proportional tax for registering a new house ownership (stamp duty).[24] This tax makes it more expensive to move, hence it will reduce the turnover rate. The correlation between mean prices and turnover would be zero – under the DGP described above – however. The tax would make it harder to find significant correlations between mean price and turnover, simply because real changes in user cost are smaller than nominal changes.

To truly gain an understanding of structural relationships, it is essential to have micro-data at the individual or household level.

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[24] The proportional tax (stamp duty) is issued upon the acquisition of owner-occupied housing and was 1.5 percent of the property value in 1999 (3 percent for a juridical person). To stimulate the housing market, the stamp duty was temporarily reduced to 0.5 percent and 1 percent, respectively, during the period June 1996 to Dec. 1997 (see Riksskatteverket, 2000).

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