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TESTING THE COINTEGRATION OF HOUSE AND STOCK PRICES IN FINLAND*

KARI TAKALA

Bank of Finland, P.O. Box 160, Rauhankatu 19, SF-00101 Helsinki, Finland

and

PEKKA PERE

The Research Institute of The Finnish Economy, Lönnrotinkatu 4 B, 00120 Helsinki, Finland

This paper looks at the efficiency of asset markets and pricing rules with regard to house and stock markets in Finland. House and stock prices are found to have unit roots, which is a necessary condition for efficient markets. However, asset prices are not pure random walks, but instead unit root processes. The unit root part is supposed to reflect the changes in fundamentals, and the error component reflects short run deviations from the market equilibrium.

Evidence on cointegration between asset prices is found. Based on the hypothesis of cointegration, an error correction model is estimated. Deviations from the long run market equilibrium can be used to improve predictions of stock and house prices, e.g. house price predictions can be improved on average by using lagged changes in the stock index and the equilibrium error as a useful indicator of disequilibrium in the cointegrated markets. The presence of such an error correction term in itself shows that asset markets are not fully efficient.

Granger causality between these two aggregate asset prices runs from the volatile stock market to the housing market rather than the opposite way. During the sample period the Finnish money market went through a gradual liberalization process that revealed an imbalance in the asset market caused partly by rationing in the rental housing market. In addition, tax incentives for owner-occupied housing and relaxed liquidity constraints significantly increased the demand for houses during the period 1987–89. Real bank lending proved to be a significant predictor for asset prices, especially for real house prices. Error correction models with additional variables due to liquidity constraints, taxation and demand for assets were used to explain the house prices.

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1. Introduction

The economic theory of capital asset pricing relies heavily on the principles of present value calculations and the hypothesis of efficient capital markets. Present value models tell us that the price of an asset is a function of the expected future yields discounted to the current date. This should apply to all assets such as stocks, land, houses and durables, since they are alternative investment objects. However, the valuations of forward-looking consumer/investors are affected by subjective factors as well, like tastes, asymmetric information, differences in time preference and degree of risk aversion.

At each moment the market demand — supply condition produces equilibrium price and quantity, which is based only on marginal amounts traded of each asset. With durable assets like houses, the prices realized on the market may not be perfect indicators of the value of the whole stock. Housing investment is also volatile and prices of new houses vary with the prices of the old stock. This may make the adjustment process slow, although prices are supposed to carry all the relevant information regarding the asset. On the other hand stock market crashes and speculative bubbles have cast serious doubts on the efficiency of financial markets in rationally reflecting the fundamentals. Therefore it might at first seem unlikely, that strong dependence would be found between markets of separate assets.

In the *short run*, asset prices vary with changes in operating cost, depreciation, alternative asset yield, taxation rules and risk premiums. However, even in the case of durable assets, changes in interest rates are not directly reflected in asset prices, as would be anticipated from the present value formula.

In the *long run*, it seems plausible that the determination of asset prices has to depend on the replacement cost of an asset, which may sometimes be hard to define for exhaustible resources — land, ground water, forest. In an efficient economy these factors of production are used to the point where the value of the marginal product equals the marginal cost of the input. Therefore, it is not surprising that the efficient market hypothesis has implications for asset prices as well. In this paper several implications of the joint hypothesis of the asset pricing theory and efficient markets are tested. The importance of these implications has recently been rediscovered, since they can be tested with the newly discovered and rapidly developing econometric methodology of unit roots and cointegration.

In the long run, asset prices should be cointegrated, and real after-tax risk adjusted returns should coincide. There is no reason to expect the prices of two different individual stocks to be cointegrated, but since we are dealing here with broad asset markets, this may be the case. This is the question we address. Unfortunately many factors can disturb the expected cointegration, like changing risk premiums, changes in expected inflation and relative prices, tax rules and tax advantages, regulation of supply, portfolio changes induced by changes in subjective time preference, irrational expectations about future returns etc. Several proxies of these variables are used in explaining the divergence of asset prices from their equilibrium values.

Quite recently *statistical models* have been developed to describe parsimoniously the choice and demand for assets. This line of work follows the mean-variance approach to asset demand, according to which the demand for an asset is fully described by two moments of the return distribution, namely the expected return and variance. This approach is taken in the strict version of capital asset pricing model (CAPM), which has thus far been the dominant model.

For example, the efficient market hypothesis applied to present value models with rational expectations implies that asset returns and hence asset prices, should follow a random walk, possibly with drift. Empirically this means that changes in asset prices cannot be predicted with current information apart from the drift. On the other hand, under efficient markets asset prices should be cointegrated, which means that their price paths should not diverge in the long run (Bossaerts, 1988). The pure random walk hypothesis and asset price cointegration are contradictory, since both cannot prevail at the same time (Granger, 1986). If cointegration prevails, this means that the equilibrium error can be used in predicting the corresponding target changes, which cannot hold under the pure random walk property.

We start our analysis with a description of the data and recent phenomenena in the development of asset prices and their determinants, in particular, the effects of the liberalization of the money market which occured in the late 1980's. Then we proceed in reviewing the implications of the CAPM in tests of prices. In the econometric section of the paper, we review the unit root and cointegration tests proposed in the literature and apply them to our data. After presenting some evidence on cointegration of asset prices, we move to the question of possible causes of short run distortions in the asset market. We try to explain the short run equilibrium error between housing prices and the general stock index with general economic fundamentals and more specific determinants of asset prices. Finally, we construct and test ECM model for house price changes, based on cointegration of asset prices with some additional explanatory factors.

2. *Description of the asset market and data*

Two basic asset series: housing prices and the general stock index for Finland were analyzed with quarterly data of 1970/1 – 1990/2. The choice of assets was motivated by their availability and overall importance as investment targets. The house price index for the whole country was used, although the houses in the Helsinki area could be more close substitute for stocks (see also Suoniemi, 1990). The stock index used is the Unitas index of the biggest commercial bank (SYP) in Finland. We have generally considered the stock market to be more efficient than the housing market because of its more rapid reactions to new information, greater volatility and smaller transaction costs. To find out whether this is true, we first review some major aspects of asset prices and their determinants from 1970.

From the nominal prices it is hard to see a close relation between the price levels, only a clear upward trend (figure 1). The real prices of houses and the stock index show already evidence of a long run linear relation, especially if we match the ranges i.e. making a linear transformation for house prices (see figures 2 and 3). The real prices were calculated by deflating nominal prices with the consumer price index (CPI). The base year in figures is 1972. During the period 1972 – 90 the real price curves intersected about four times. The dramatic increase since 1986 in stock index and, a little later, housing prices

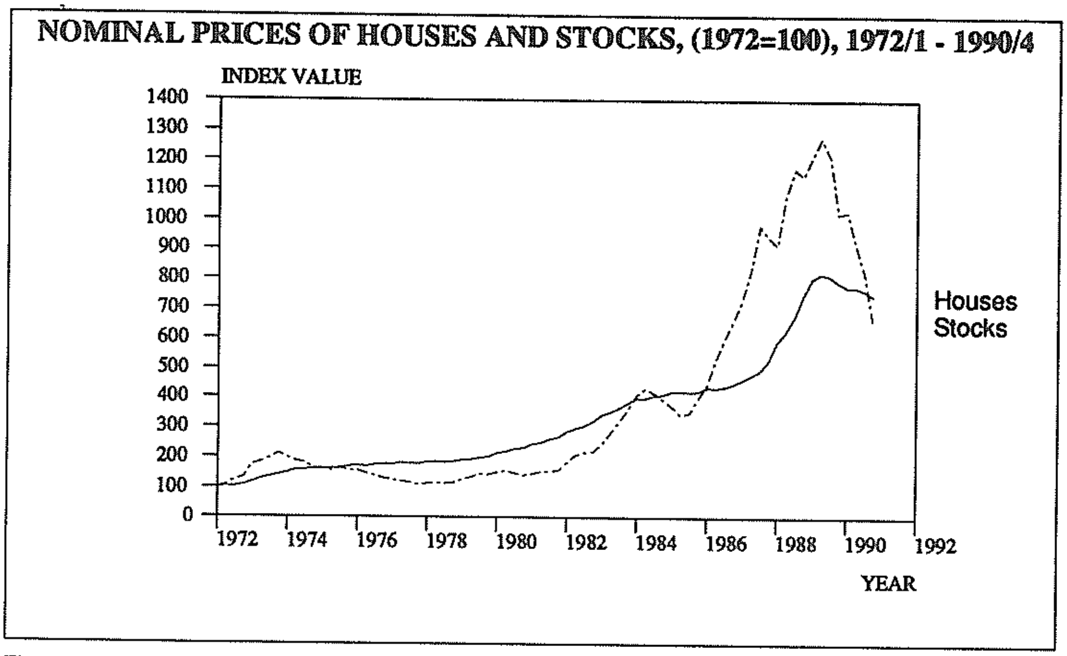
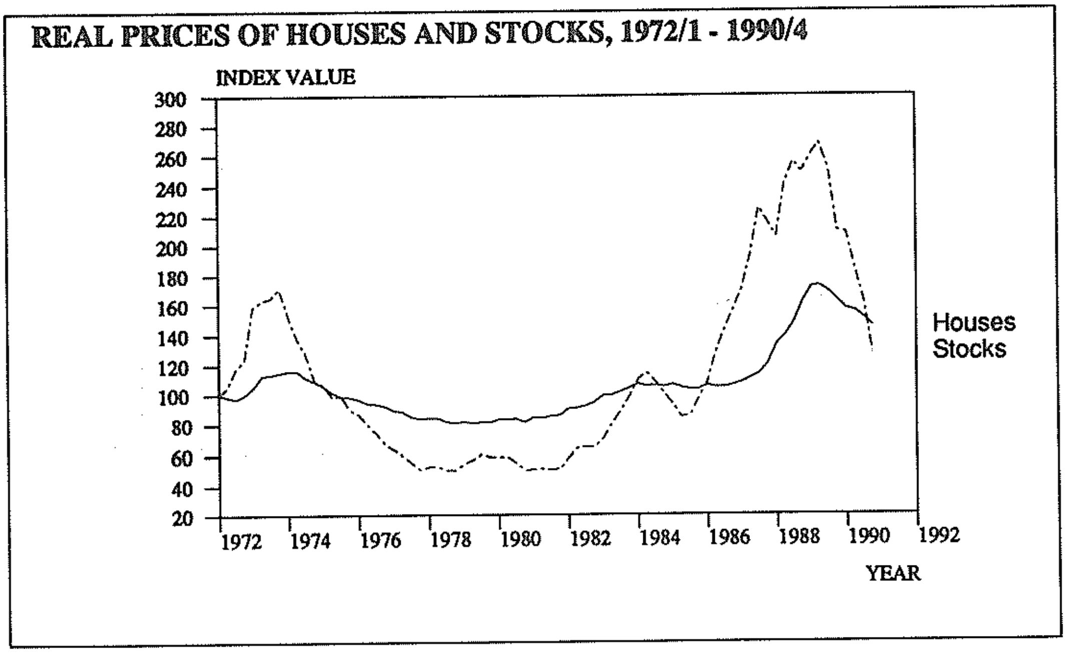


Figure 1.



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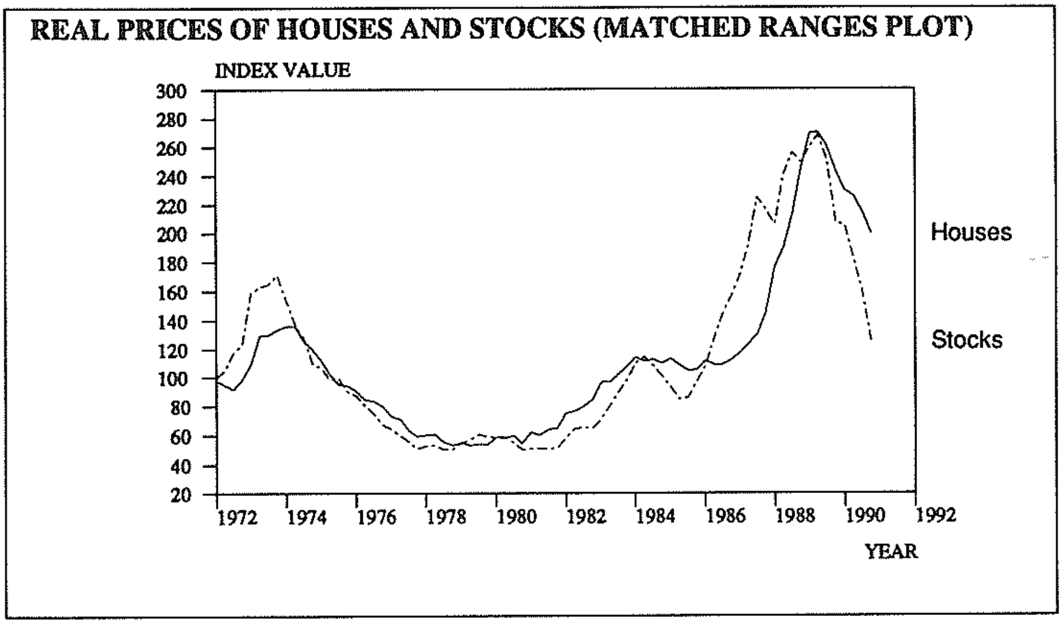


Figure 3.

was mainly a consequence of capital market liberalization and increased bank lending thereafter. The October 1987 stock market crash was not very deep in Finland, and the level of general stock prices recovered to the pre-crash level already in spring 1988. From spring 1989 real price indexes have fallen as a consequence for a similar overheating like in Norway, West-Germany and Sweden.

The relation between real bank lending and

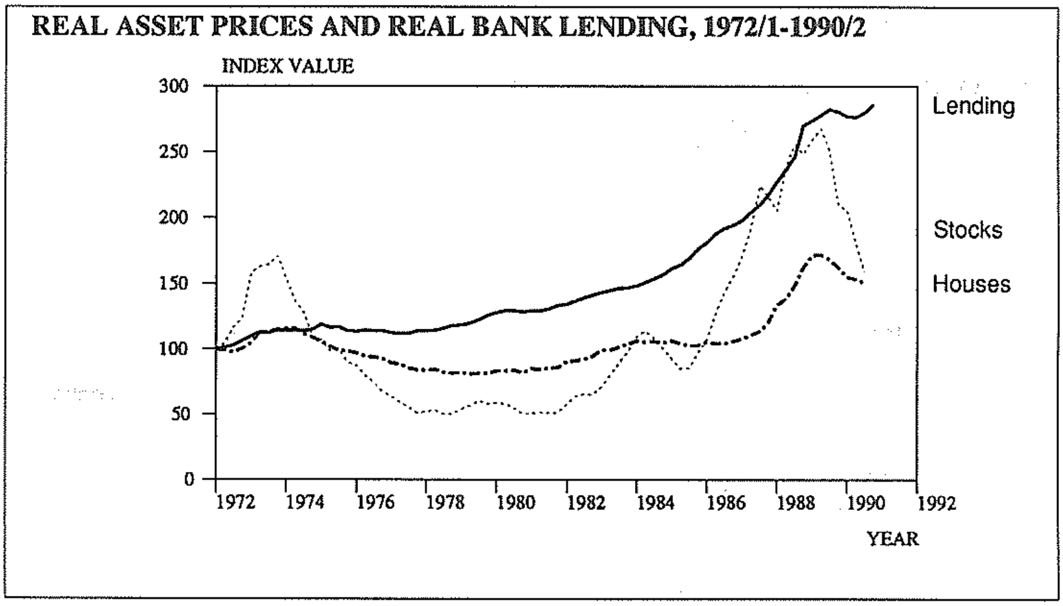


Figure 4.

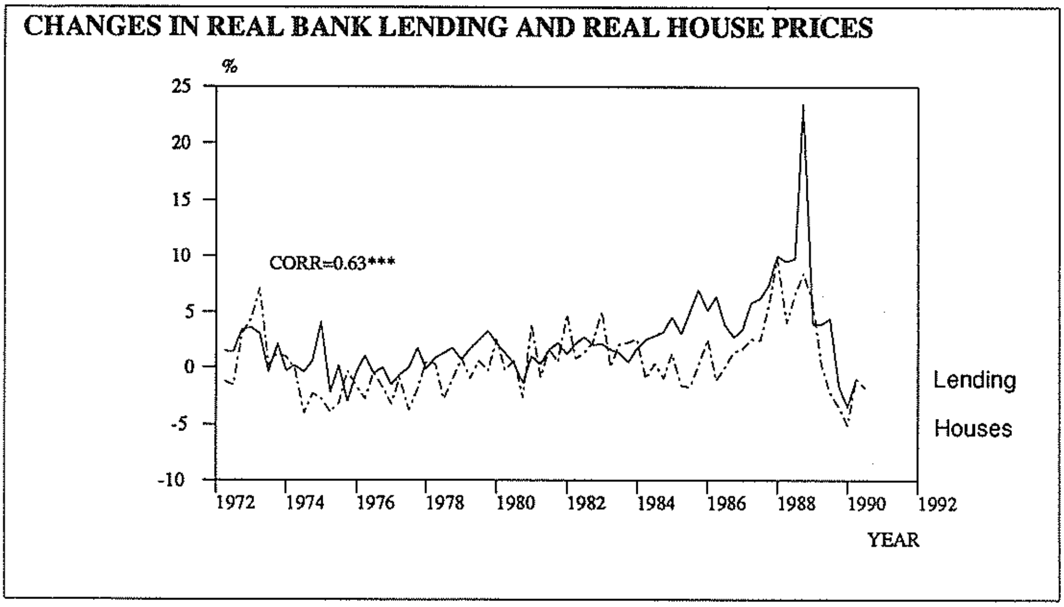
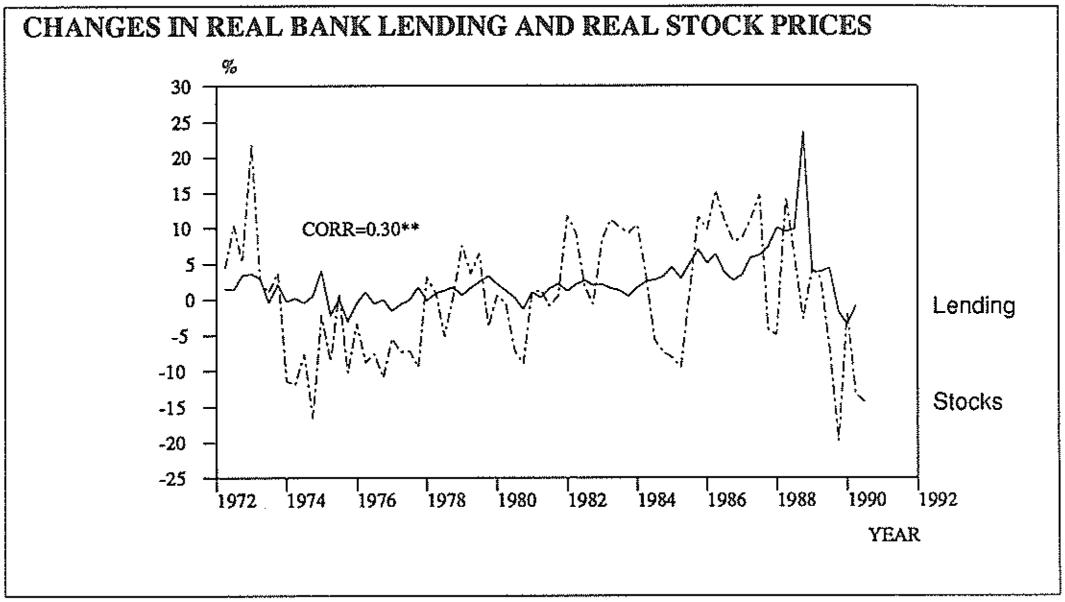


Figure 5.

real house prices, deflated with the consumer price index, can be seen from figure 4. The differences of real bank lending and asset prices can be seen from figures 5 and 6. The instantaneous correlation between real bank lending and real house price difference was even as high as 0.63. The liberalization of the money market revealed the imbalance in the rationed rental housing market. The very drastic symptom of this was the large demand in-



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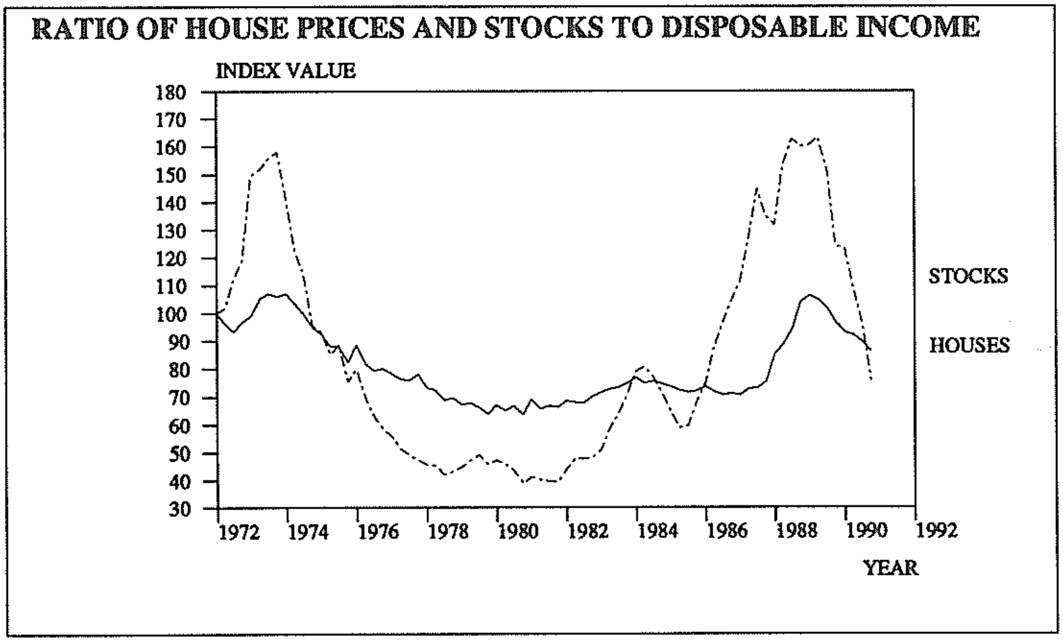

Figure 7.

crease for owner-occupied housing. The increase in demand and bank lending went directly into asset prices, since supply was almost inelastic. The Finnish rental housing market has been rationed since 1960's and tax incentives for owner-occupied housing have risen the share of owner-occupied housing since then from 56 % up to over 70 %.

The ratios between nominal house prices and nominal stock index to household dis-

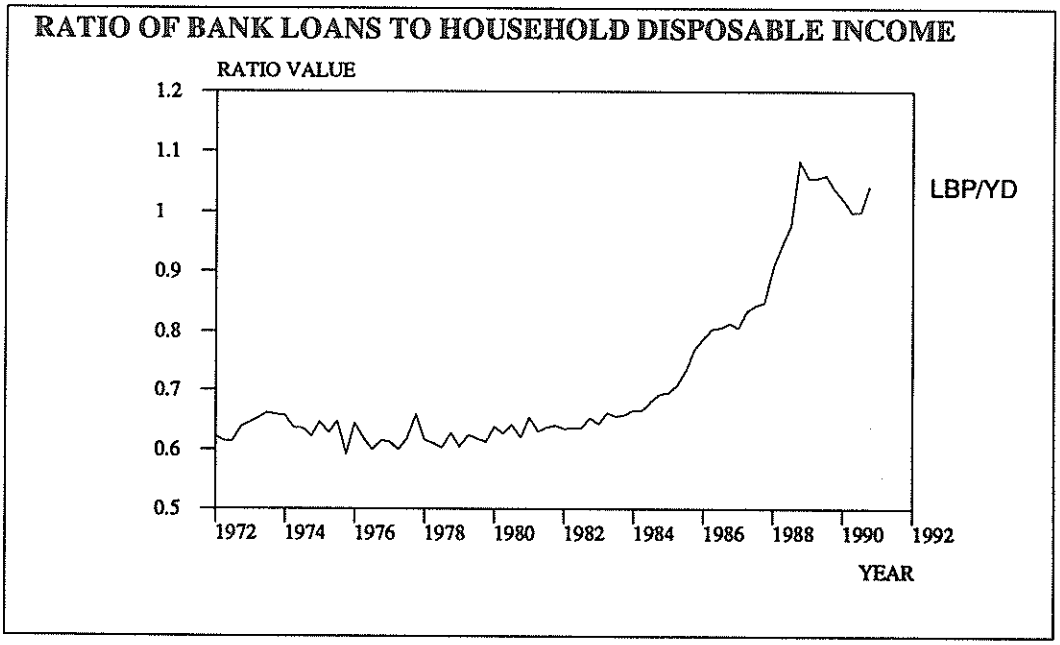

Figure 8.

posable income can be seen from figure 7. In the beginning of 1970's house prices went up following a demand boom. After the 1974 oil crisis, house prices decreased steadily in real terms with respect to income till the end of 1970's. Real construction costs have been quite stable and construction costs and consumer prices increased at about the same growth rates up to 1987.[1]

The liberalization of capital markets and foreign capital inflow through firms and banks (households are not yet able to raise foreign loans directly) is the primary candidate to explain the rapid stock and house price rises since 1986. The share of housing loans in total bank lending has however fallen slightly, as consumption loans have risen even faster. Because the value of the housing stock has risen rapidly, the ratio of housing loans to the value of housing stock has increased only slightly. Nevertheless, it is clear that the debt/asset ratio has been increasing. Figure 8 presents the ratio of outstanding bank loans to household disposable income. The planning horizons of households may have lengthened due to money market liberalization, since liquidity constraints have eased and borrowing against future earnings is easier.

3. *Implications of capital asset pricing model*

The aim of asset pricing theories is to explain why different assets have different risk premiums. Therefore, the purpose of the capital asset pricing model is to show how expected returns of portfolios are also functions of the riskiness of assets (Rothschild, 1985). The riskiness of an asset is usually defined as the variance of the distribution of the return. The investor's choice of a portfolio is based on risk adjusted expected returns. The core of portfolio theories is that the risk of an asset depends on the return of other assets as well. In this respect CAPM relies on the concept of

[1] *It has been suggested by the construction firms that house prices increase, because construction cost rise. This is not quite true, since change in the prices of newly built houses follow changes in the prices of the existing stock, which depends on the overall demand — supply situation. Therefore, the causality runs more likely from housing prices to construction costs (wages, materials) and land prices. Tobin's q approach seems to work quite well for the Finnish housing market (see Takala and Tuomala, 1990).*

efficient markets. In fact, in an efficient market, the return on an asset includes a premium for risk that cannot be removed through diversification. More importantly the risk premium is proportionate to the covariance between asset and market returns, not to the variance of the asset itself.

Capital asset pricing theory has many applications in economic modelling. Present value models are often used in assessing the profitability of investment projects, but they can be applied as well to the valuing the price of assets — houses, land, firms and durables. The standard version of the CAPM assumes a number of requirements, like returns should be normally distributed, borrowing and lending possible at risk-free rate and there exist no transaction costs or taxes, which restricts the scope of the theory.

One important implication of the CAPM is that the observed risk premia should be fully explained by the covariance between asset returns and market return. However, in practice we have acknowledged evidence on incompleteness and inefficiency in capital markets and there is substantial evidence on the volatility clustering of asset price changes. There is no direct reason for risk or risk premiums to be constant over time. Still, risk is the main determinant of asset price changes. The varying risk premia over time is due to unanticipated interest rate movements. This risk can be measured by the conditional variance of an asset return. Another implication of the CAPM under constant absolute risk aversion is that the observed mean return and variance of return should be positively correlated across assets (Engle, Lilien and Robins, 1987 p. 394). To take into account the variation in risk, we included the conditional variance of the stock index model into the error correction models.[2]

We do not give any full representation of the CAPM model here. We merely list some of the possible ways to test implications of the theory. For a more extent exposition see Elton & Gruber (1987) or Rothschild (1985). In the empirical part, we measure the return of an asset by changes in the asset price as an implication of the present value model (see Campbell and Shiller, 1988).[3]

Testable implications:

1) The CAPM is based on the idea that all investors choose mean-variance efficient portfolios. Expected return on a risky asset is therefore proportional to the non-diversifiable risk in an efficient market. Risk is measured by the covariance of asset return to market portfolio return. Market portfolio is defined to be the linear combination of all assets, that satisfy the efficient mean-variance condition. Therefore CAPM implies that the asset risk premium is determined by the diversifiable risk, which is measured by the covariance with the market returns and which is called beta, in reference to the regression slope (Bollerslev et. al., 1988 p. 116).

The expected return (change in the asset price) of an asset i is

$$(3.1)\quad E(r_i) = r_f + \beta_i [E(r_m) - r_f],$$

where $\beta_i = Cov(r_i, r_m)/Var(r_m)$ and r_f = return on risk-free asset r_m = return on market port-

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[2] *As the CAPM is based on the idea that asset portfolios are held as functions of expected means and variances of the rates of return, shifts in asset demand imply changes in both of these determinants (Engle, 1982, p. 989). The standard CAPM should be a good approximation to asset pricing when the marginal utility of consumption is highly correlated with the return on the stock market, or more generally with the portfolio of tradable assets. Since human capital is not fully tradable, relaxing the liquidity constraints should make assets more sensitive to interest rates and return. The efficiency of the asset market or the rationality of agent behaviour can be tested by calculating the correlation between the market portfolio return and the marginal utility of consumption. The consumption CAPM implies that the returns on market portfolio and marginal utility of consumption should be perfectly negatively correlated (Blanchard and Fischer, 1988 p. 509).*

[3] *The present value model expresses the price of an asset as a deterministic function of expected future yield. With finite time yield can be separated into capital income and appreciation. Here, we assumed that asset is held forever and change in asset price includes all the changes in the discounted future capital income (dividends or imputed rent on housing). All the relevant information about yield is contained in the expectation of an asset price, since the spread between the true asset price and its expectation is stationary. If the yield follows a random walk, so would the asset price, when discount rate is constant. Campbell and Shiller (1988) consider the conditions, when yield and asset price are cointegrated.*

folio. In other words, we have the usual expression for efficient portfolios

Expected return = risk free rate
+ risk premium.

Risk premium is proportional to beta, which reflects the sensitivity of an individual asset to movements in the market portfolio return. Thus the CAPM predicts that investors in security i will be rewarded only for market related risk.

2) Engle et al. (1987) suggest that the degree of uncertainty in asset returns varies over time and any increase in the expected return can be identified from the risk premium. Investment decisions are always based on expectation, conditional on some undefined information set. We may restrict the information set to current and past values of returns, but it is unlikely that asset price processes will be stationary over time periods long enough to estimate expected returns with any accuracy. The constancy of risk premium and the variability in β can be tested with iterative least squares estimation in asset equation above.

If in the CAPM, the regression coefficient β changes over time, we are faced with the problem of non-constant conditional variance of the error term. This can be analyzed by means of ARCH models. One period rate of return should be ex ante unforecastable in mean, i.e. asset price levels are martingales, but if an ARCH-M model can be identified, economic gain could be achieved through changes in diversification (Engle, 1988).

Portfolio theory emphasizes the interdependence between the risk of particular asset and the risk of other available assets. The riskiness of an asset depends ex post on how much the actual return differs from the expected return. For a riskier asset the required rate of return is higher because of the higher risk premium. But over the efficient market risk, adjusted returns should be equal. Therefore different observed nominal returns simply reflect different expectations of risk.

Since subjective expectations cannot be directly observed from market data, we may assume that rational economic agents do not make forecasting errors on average in assessing this risk. Hence at the market level, expected risk is equal to observed risk. But we cannot assume that the variances of expected and actual returns coincide. We may assume that the variance of the expected return is smaller than the actual variance, since predictions on return are usually based on a smaller information set. Predictions are made for the systematic part of the actual variation and since there is always some positive error variance, predictions are conservative. In addition to information processing and analyzing, the econometrian's information set is smaller than that of the market participants.

We sum up the hypotheses to be tested as follows;

a. Random walk property and unit roots

H_o: Asset returns follow a random walk, and asset prices a random walk with drift[4]

Asset returns should follow random walk process under informational efficiency; therefore,

(3.2) $r_t = r_{t-1} + e_t$, where $e_t \sim NID(0, \sigma^2)$.

Asset prices should be random walk with drift (μ), which can be tested with

$$H_o\colon\ r_t = \mu + r_{t-1} + e_t \quad \text{against}$$
$$H_1\colon\ r_t = \mu + bt + \phi r_{t-1} + e_t,$$

where b is the slope of the deterministic time trend and $\phi < 1$. The appreciation of asset prices were used as a proxy for the returns, although dividends is a form of return for stocks in the short run as well.

b. Cointegration and Granger causality

H_o: CAPM implies the cointegration of asset prices, where all risky assets are held in the same proportion, in addition risk adjusted returns should coincide (in the long-run) with the return of a riskless asset (Rothschild, 1985).

Ordinary asset pricing models like the CAPM imply complete separation, which

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[4] *Exceptions are also possible. For instance, the Lucas (1978) asset pricing model implies that stock prices do not have to have the martingale property, except under risk neutrality (Michener, 1982 p. 166).*

makes asset prices perfectly collinear. However this result can usually be rejected with empirical data. If we relax the assumption about perfect separability and replace it by the weaker separating equilibrium, we may preserve approximate collinearity (Bossaerts, 1988). This leads to cointegrated asset prices, which allows for weakly dependent pricing errors. The rise of price of an asset will lead to an increase in the portfolio share of this asset, but the ratio $p_{it}*s_i/p_{jt}*s_j = w_{ij}$ should be time-invariant, where s is supply of asset and i, j denote the assets. Unfortunately supply for an asset is rarely known in practice. For housing the supply may be approximated with current stock plus housing investment, but for stocks new share issues are very flexible and it may be hard to define supply.

H_0: cointegration implies that the stationary equilibrium error term should Granger-cause at least one of the cointegrated variables (Engle & Granger, 1987, Campbell and Shiller, 1988 p. 507).

For finding Granger's predictive causality between variables, non-cointegration should be checked first. If cointegration exists, the equilibrium error should be added as an additional regressor to the causality tests of stationary variables.

4. Econometric methodology

We use the increasingly popular cointegration methodology. In this section we review very briefly the concepts of integration and cointegration and the related tests that we use. See for example Diebold and Nerlove (1988) and Pere (1990) for more elaborate surveys.[5]

The importance of cointegration is due to the empirical finding that many time series resemble I(1) time series and the fact that economic theory often suggests that the levels of the time series should be closely connected. Thus we see that many economic time series should be CI(1, 1). In addition, according to the so called Granger representation theorem [Engle and Granger (1987)] vector x_t has an error correction representation, if the variables are cointegrated. So we have a handy way to model the series, if the data tell us that the series are cointegrated.

We use these concepts to test if stock and house prices are cointegrated. First we explore the order of integration of these series. We use the Dickey-Fuller -tests (Dickey and Fuller (1979)) for this purpose. Their idea is very simple. Let us assume that x_t is an AR(1) process

$$(4.1) \quad x_t = \phi x_{t-1} + u_t,$$

where $u_t \sim NID(0, \sigma^2)$. We can reparametrize (adding $-x_{t-1}$ on both sides) this as

$$(4.2) \quad Dx_t = (\phi - 1)x_t + u_t.$$

This equation could be expanded by adding seasonal terms and linear time trend

$$(4.3) \quad Dx_t = \mu + bt + \sum_{s=1}^{3} \Phi_s Q_{st} + (\phi - 1)x_t + u_t.$$

Our null hypothesis is that x_t is I(1) or $\phi = 1$. If the null is correct, then in the regression (4.2) the estimate of $r = (\phi - 1)$ and its t-value should be near zero. If the t-value differs significantly from zero, we reject the null hypothesis and conclude that $r < 0$ or $\phi < 1$ or equivalently that x_t is stationary. Otherwise we accept the null hypothesis. This is the Dickey-Fuller test (DF).[6]

If it turns out that our data has to be modeled with an AR(p) autoregression, where $p > 1$, to whiten the residuals u_t:

$$(4.4) \quad x_t = \sum_{i=1}^{p} \phi_i x_{t-i} + u_t.$$

[5] *A time series x_t is said to be integrated of order d, if it has a stationary and invertible ARMA representation after differencing d-times. If this is so, we denote $x_t \sim I(d)$. Many economic time series are supposed to be I(1) or some times I(2) series, see for example Nelson and Plosser (1982) and Hall (1986).*

The elements x_{it} of a N-variate time series x_t are cointegrated of order d, b if (i) each $x_i \sim$ I(d), i = 1, . . . ,N, and (ii) there exists a vector β ($\neq 0$) such that $\beta' x_t \sim I(d-b)$. We denote this by $x_t \sim CI(d, b)$. That is time series are cointegrated, if they are integrated of the same order and some linear combination of them is integrated of a lower order than the original series.

[6] *In our regression here the appropriate distribution theory is nonstandard because of the nonstationarity of x_t. However, we can use the t-value as a test statistic, because the textbook of Fuller (1976) includes its critical values as simulated by Dickey.*

This again can be reparametrized to

$$(4.5)\quad Dx_t = r_1 x_{t-1} + \sum_{i=2}^{p} r_i\, Dx_{t-i+1} + u_t,$$

where $r_1 = (\Sigma_{i=1}^{p} \phi_i - 1)$ and $r_j = -\Sigma_{i=j}^{p} \phi_i$, $j = 2,\ldots,p$. We can still test the hypothesis null $x_t \sim I(1)$ with the help of the Dickey-Fuller test. The null is easily seen to be equivalent to $r_1 = 0$. The augmented Dickey-Fuller test (ADF) statistic is the t-value of r_1. The test is otherwise exactly the same as the ordinary Dickey-Fuller test. The critical values in Fuller (1976) apply to this statistic also, if the sample is big enough. If we accept the null hypothesis, we go on to test for cointegration.

We use these same tests to explore the cointegration of our data. After checking that house prices (y_t) and stock prices (x_t) are I(1) we estimate the so called cointegrating regression

$$(4.6)\quad y_t = ax_t + u_t.$$

Now if y_t and x_t are to be cointegrated u_t has to be stationary. We can test this by applying the DF and ADF test to u_t, now called DFR or ADFR respectively. According to the null y_t and x_t are not cointegrated or u_t is integrated. The alternative is that y_t and x_t are cointegrated or u_t is stationary. The test goes exactly as the ordinary DF and ADF tests except that we cannot use the tables in Fuller (1976), because the residuals $\hat{u}_t$ are estimated. Due to the properties of ordinary least squares, the estimated residuals look more stationary than the original residuals and this changes the critical values of the DF and ADF tests. Engle and Granger (1987) have simulated some critical values. We use them as approximate critical values for our time series, which has become a common practice in empirical studies of cointegration. Small sample critical values for ADF test can be found in Blangiewicz & Charemza (1990).

We report also the Durbin-Watson statistic (CRDW) of the regression (4.6). It is a quick check of the first order of autocorrelation of u_t, because CRDW $\approx 2(1-\hat{r})$, where $\hat{r}$ is the first order sample autocorrelation of u_t. This gives us some indication of the probability of cointegration of y_t and x_t. DW is usually only slightly above zero if y_t and x_t are not cointegrated. If u_t were an AR(1) process, it could be used as a test for cointegration (see Engle and Granger (1987)). Usually, as in our data, this is not the case, so we don't refer to it as a proper test statistic. In addition the Johansen (1988, 1989) tests for cointegration vectors, available in PC-GIVE, were calculated.

The implications for Granger causality in cointegrated data could be done as follows. If y_t is the housing price index and x_t stock index, then

$$(4.7)\quad z_t = [a_1\ a_2][x_t\ y_t]' = a_1 x_t + a_2 y_t$$

The proposition of one-sided Granger causality from equilibrium error term z_t to Dx_t can be seen from the regression

$$(4.8)\quad Dy_t = \mu + \sum_{j=1}^{p} b_j\, Dy_{t-j} + \sum_{j=1}^{p} w_j\, Dx_{t-j} + \sum_{j=1}^{p} q_j\, z_{t-j} + e_t,$$

where the null hypothesis for Granger noncausality is H_o: $q_j = 0$, for all j. The Granger causality implication is a consequence of the expectation theory, namely that the price of an asset is a weighted sum of expectations of the future capital income and appreciation.

We want to emphasize that the results of these various tests are tentative. The tests are known to have low power against very autocorrelated but still stationary alternatives (Dickey and Fuller (1979) and Engle and Granger (1987)). Further, the test size may be distorted by changes in the process generating the data (Hendry ja Neale (1990)) and MA-terms in the series (Schwert (1989)). And last but not least there is the question of the quality of the house and stock indexes. With these caveats in mind let us proceed to the empirical analysis.

5. *Empirical results*

We begin with a description of the unit root and cointegration characteristics of the series under consideration. Then several Granger causality results are given, after which an estimated ECM is presented, along with its implications regarding equilibrium error.

5.1 Results from unit root tests

The sizes of the AR(1) coefficients for the price series immediately suggested the presence of unit roots in logaritmic levels (table 1). The DF and ADF tests confirmed this conclusion. Despite this, the residuals of these autoregressions were clearly autocorrelated. Therefore prices are not pure random walk processes or martingales. In the augmented Dickey-Fuller test five lags of the dependent variable were added to the equation to eliminate autocorrelation in the residual.

Table 1 presents different versions of unit root tests with constant linear time trend and seasonal dummy variables for house and stock prices. Both DF and ADF tests showed no evidence against unit roots in the price level series. For log differences of the asset prices unit roots are rejected. Surprisingly both asset price series included significant seasonal variation. For the stock index the first

Table 1. Unit root tests for asset prices;
Dickey-Fuller and Augmented Dickey-Fuller tests
Sample period: 1970/1 – 1990/2, 82 observations

Variable	RLPHM	RLUNI	DLRPHM	DLRUNI
Dickey-Fuller regressions:				
$x_t = \mu + \phi x_{t-1} + u_t$				
Reg.coeff. (ϕ)	1.02	0.99	0.54	0.54
DF -test	0.83	−0.32	−4.80***	−4.70***
Order of integration	I(1)	I(1)	I(1)	I(1)
$x_t = \mu + bt + \sum_{s=1}^{3} \Phi_s Q_{st} + \phi x_{t-1} + u_t$				
Reg.coeff. (ϕ)	1.00	0.99	0.51	0.54
b (t -test)	1.99*	1.07	1.17	0.29
Q_s (t -tests for seasonals)	—	2.17 Q_1	−2.90 Q_2	3.67 Q_1
DF -test	−0.26	−0.73	−4.95***	−4.67***
Order of integration	I(1)	I(1)	I(0)	I(0)

Variable	RLPHM	RLUNI	DLRPHM	DLRUNI
Augmented Dickey-Fuller regressions:				
$x_t = \mu + \phi x_{t-1} + \sum_{j=1}^{5} g_j Dx_{t-j} + u_t$				
Reg.coeff. (ϕ)	0.98	0.98	0.55	0.53
ADF -test	−0.92	−1.28	−3.40**	−3.19**
Order of integration	I(1)	I(1)	I(0)	I(0)
$x_t = \mu + \phi x_{t-1} + bt + \sum_{s=1}^{3} \Phi_s Q_{st} + \sum_{j=1}^{5} g_j Dx_{t-j} + u_t$				
Reg.coeff. (ϕ)	0.98	0.97	0.51	0.55
b (t -test)	1.21	0.74	1.34	0.30
Q_s (t -tests for seasonals)	—	3.03 Q_1	−2.02 Q_2	3.01 Q_1
ADF -test	−1.01	−1.45	−3.33**	−3.02**
Order of integration	I(1)	I(1)	I(0)	I(0)

Abbreviations in prefixes of variables are L = logarithmic and D = difference, R for real (deflated with the consumer price index). PHM refers to the index of housing prices in the whole Finland and UNI to the UNITAS stock index of the Helsinki stock exchange.

quarter of the year has the highest value; in house prices, change slows down or becomes negative in the second quarter. Except for the real house price level in the DF test, there was no indication of deterministic time trend.

The logarithmic transformation dampened the exponential rise in asset prices during the financial market liberalization period. Since our estimation period covered only 20 years, we cannot be sure whether this simply indicates an exceptional period. The long upward swing of stock and house prices in the 1980's after the oil crises was in any case an exceptional period. Some part of it dues to increase in money supply throughout the western world. It seems however, more reasonable that asset prices follow a stochastic rather than a deterministic time trend.

5.2 Cointegration tests

Results of the cointegration tests (CRDW, DFR and ADFR) are shown in table 2. For a quick check the CRDW test between the price of houses and stock price index levels gave no indication of cointegration. Although

Table 2. Cointegration tests for asset prices;
Cointegration Durbin-Watson, Dickey-Fuller and Augmented Dickey-Fuller tests,
Sample period: 1970/1 – 1990/2

Tests:

CR: $y_t = ax_t + u_t$
DFR: $Du_t = -ru_{t-1} + e_t$
ADFR: $Du_t = -ru_{t-1} + \Sigma b_j Du_{t-j} + e_t$, $j = 1, \ldots, 4$.

Cointegrating Regression: $y_t = ax_t + u_t$

Example: LRPHM = 3.00 + 0.34 LRUNI, $R^2 = 0.84$, DW = 0.18
(36.7) (20.8)

Dependend variable:	RPHM	RUNI	LRPHM	LRUNI
CRDW	0.23	0.25	0.18	0.18
Dickey-Fuller Regression $Du_t = -ru_{t-1} + e_t$				
Reg.coeff. (−r)	−0.08	−0.10	−0.06	−0.08
DFR	−1.37	−1.83	−1.18	−1.58
Augmented Dickey-Fuller Regression: $Du_t = -ru_{t-1} + \sum_{j=1}^{4} b_j Du_{t-j} + e_t$				
Reg.coeff. (−r)	−0.30	−0.26	−0.25	−0.22
ADFR	−4.29***	−4.22***	−3.94***	−3.80***

The critical values for 100 observations are presented in Engle & Granger (1987) and for smaller samples in Blangiewicz & Charemza (1990, p. 306).

The Johansen (1988, 1989) tests of the rank of cointegration
Tested variables: LRPHM, LRUNI

H (rank)	Eigenvalue	Maximal eigenvalue		Trace test	Crit.value
		Test	Crit.value		
	μ_i	$-T \ln(1-\mu_i)$	$\mu_{max}(.95)$	$-T\ \Sigma \ln(1-\mu_i)$	$\mu_{trace}(.95)$
$r \leq 1$	0.035	2.72	8.08	2.72	8.08
$r = 0$	0.218	18.98	14.60	21.69	17.84

the DW statistic in cointegration regression is smaller than R^2, this is not a case of spurious regression, but merely a result of the fact that regression and correlation measures the co-movement of variables. Cointegrating regressions without R^2 close to unity should be viewed with caution. Our R^2 on regressing logs of real house prices on corresponding stock index was 0.84.

However, taking residual autocorrelation into account inference, the ADFR test with 4 lags, results in strong rejection of non-cointegration between both real asset prices and their logs. This rejection, i.e. the acceptance of cointegration, was significant with 1 % level in this sample. Also the Johansen (1988, 1989) type of cointegration tests accept cointegration between asset prices with 5 % significance level (table 2).[7]

It seems highly probable that cointegration would be a useful working hypothesis. The relation between asset prices can be interpreted as an equilibrium process, that takes about 1–3 quarters to adjust to a smaller shock in the asset market. Since the equilibrium mechanism seems to be working rather slowly, the error correction term is nearly non-stationary.

Table 3. Granger causality tests for asset prices.

Dependent Dy_t	Explanatory Dx_t	F -test for restriction[1] without null lag	with null lag
Changes in real asset prices			
DLRPHM	DLRUNI	2.11*	1.84
DLRUNI	DLRPHM	1.78	1.56
Real asset prices and real lending			
DLRPHM	DLRLBP	2.72*	5.51**
DLRLBP	DLRPHM	1.76	4.52**
DLRUNI	DLRLBP	1.59	2.16
DLRLBP	DLRUNI	1.91	2.43*
Real asset prices and error correction term			
DLRPHM	ECM	3.45**	12.21**
DLRUNI	ECM	1.77	10.30**
House prices and interest rates			
DLRPHM	RLB	1.31	2.03
DLRPHM	RB	0.39	1.40
DLRPHM	DRLB	1.38	1.78
DLRPHM	DRB	1.90	2.73*
Stock prices and interest rates			
DLRUNI	RLB	1.15	1.57
DLRUNI	RB	2.47*	2.07
DLRUNI	DRLB	1.89	2.05
DLRUNI	DRB	2.44*	2.09
House prices and construction costs			
DLRPHM	DLRPIH	0.35	0.47
DLRPIH	DLRPHM	2.60*	2.34*

[1] Granger causality tests were performed as bivariate autoregressive-distributed lag models (max lag used was 5) as

$$Dy_t = \alpha_0 + \sum_{i=1}^{5} \alpha_i Dy_{t-i} + \sum_{j=0,1}^{5} \beta_j Dx_{t-j} + h z_{t-1} + e_t$$

with quarterly data. The results for null hypothesis (Granger non-causality from x_t to y_t) H_0: $h = \beta_0 = \beta_1 = \ldots = \beta_5 = 0$ can be seen above (see Hendry, 1989 p. 36). F -statistic had degrees of freedom F(5,65) without null lag and F(6,64) with null lag, when cointegration term was not present. With cointegrated variables (y_t, x_t), the error correction term (z_{t-1}) also has to be included in the regressions (see Granger, 1988).

5.3 Granger causality tests

Granger causality (GC) tests were performed for stationary dependent and explanatory variables in a single equation context, since strong exogeneity of the regressors with respect to the parameters of interest is based on the assumption of Granger non-causality from past and present values of the dependent variable. With cointegrated variables this predictive causality can also run through the error correction term, which could be invalidated unless the error correction term is taken into account. The results of the GC tests are presented in table 3.

From visual inspection of the asset prices and Granger causality tests it seems probable that causality runs from the stock index to the house prices rather than vice versa (figure 2).

[7] Unit root and cointegration tests were also performed with annual data covering 1965–1989 (25 observations). The only house price data (namely bank financed houses in the Helsinki area) available for this period was found in a study of C. Bengs (1989). However, the cointegration tests were quite similar for these series. The CRDW test result was just below the critical rejecting non-cointegration level (DW = 0.32 with LPHM regressed on LUNI), but the DFR and ADFR tests indicated rejection of non-cointegration.

This conclusion is further evidenced by the fact that the error correction term Granger causes house prices even at the 1 % significance level. This interpretation accords with the usual assumption of the stock market reacting rapidly in changes in economic fundamentals. House prices seem to be more rigid.

Our hypothesis on the strong impact of real bank lending on real house prices was supported. The instantaneous correlation between real bank lending and real house prices was 0.63 and with lending lagged one period 0.51. Bank lending Granger causes real house prices at a very high significance level. Further, the causality seems to be one-sided, namely increases in lending increase prices, but high prices do not predict high lending. The effect of real bank lending is not so strong in the stock market, although the instantaneous correlation was still significant (see figures 5–6). We must emphasize that real bank lending is merely an useful variable in predicting house prices, not a true causal factor itself.

No Granger causality was found between real logarithmic disposable income and real asset prices. Interest rates and house prices were not significantly correlated. Government bonds and stock shares could be substitutes, since the former predicts opposite changes in latter. One interesting relation was found between housing prices and construction costs. It appears that increasing house prices predict increases in costs rather, than that rising construction costs being the main reason for rises in housing prices.

5.4 Error correction model for house prices

In table 4 an ECM is estimated for the change in housing prices. There was significant autocorrelation in the dependent variable that was handled with a lag in the dependent variable. The sign of the lagged error correction term was negative as expected from ECM -theory. Table 4 also gives the diagnostics summarizing the adequacy of the formulation. Clearly this basic version of the model is not adequate enough. Problems emerge with respect to autocorrelation of the residuals, heteroscedasticity and misspecification of the functional form.

In table 5 a more elaborated version is presented. In addition to the basic variables, several constructed variables were used to describe the ongoing changes in the asset market due to liberalization of financial market. These additional variables were always lagged at least one quarter (see appendix for constructed variables). A liquidity constraint proxy was constructed on the basis of the difference between the market interest rate (RS) and the bank lending rate (RLB) multi-

Table 4. Error correction model for house prices
Quarterly data, 1970/3–1990/2
Modelling DLPHM by OLS

Variable	Coefficient	STD error	t-value	Partial r^2
DLRPHM 1	.370	.110	3.367	.132
DLRUNI 1	.071	.035	2.013	.051
ECM 1	−.074	.035	−2.090	.055
CONSTANT	.001	.004	.327	.001
Q2	−.018	.006	−3.023	.109
DLRLBP 1	.397	.192	2.064	.054

Model performance		Diagnostics for the residuals	
R^2	0.53	DW	2.24
R^2 (diff.+seasonal)	0.37	AR(5); F(5.69)	2.10
F(5,74)	16.93	ARCH; F(4,66)	0.10
		Normality X^2 (2)	2.59
		Heterosced. F(9,64)	0.55
		Functional form F(19,54)	3.70**
		Reset F(2,73)	4.77*

plied by the ratio of bank loans to disposable income (LBP/YD). The log value of consumption was used as a proxy for permanent income, since it is a result of optimizing behaviour in light of all relevant information on the budget constraint available to consumer/ investors. Innovations in the AR(1) model of disposable income differences were used to account for surprises in income. In addition a couple of tax advantage variables (absolute and relative) were constructed in order to capture the benefit from the tax deductibility of housing loan interest payments and tax exemption of imputed capital income from owner-occupied housing. These variables were based on a estimate of the mean marginal tax rate calculated in the BOF4 -model (Tarkka et.al., 1989).

The autoregressive conditional error variance of the stock index (UNI-ARCH) was used as a proxy for the varying risk premium. Risk premium is paid to the asset as a compensation for the variance in the return. There is evidence of volatility clustering in the Finnish stock market, which can be used in the house price equation as well. This term got a significant (although in many variants unstable) t -value. Under the CAPM an increase in conditional variance of the stock index could indicate increasing uncertainty that has a positive effect on housing prices and therefore a increasing yield of housing. Since housing and stock prices are clearly strongly positively correlated, we infer that increasing volatility in the stock market reduces the asset prices.

5.5 Liberalization of the financial market

During our sample period two phases of overheating in housing prices occured. Both of these phases, 1972–73 and 1987–88, were due to changes in house financing. Finland has gone through the same sort of gradual money market liberalization as the other Scandinavian countries. Interest rate constraints on bank lending to the public were abolished in steps from 1983 to August 1986. In spring 1987 free money market interest rates were allowed to be used as reference rates for non-housing loans. In addition, firms have been free to engage in foreign capital transactions.

Table 5. Error correction model for house prices
Quarterly data, 1970/3–1990/2
Modelling DLPHM by OLS

Variable	Coefficient	STD error	t-value	Partial r^2
DLRPHM 1	.253	.101	2.493	.0826
DLRUNI 1	.077	.033	2.276	.0699
ECM 1	–.069	.033	–2.102	.0602
CONSTANT	.0006	.005	.111	.0002
Q2	–.0135	.005	–2.351	.0742
D88/1	.0610	.020	3.023	.1170
DLRLBP 1	.4724	.181	2.601	.0893
LIQCON 2	–.0006	.000	–1.976	.0536
DLRCV 2	–.3918	.152	–2.562	.0869
UNI-ARC1	.0090	.004	2.251	.0685

Model performance		Diagnostics for the residuals	
R^2	0.66	DW	2.19
R^2 (diff.+seasonal)	0.54	AR(5); F(5,69)	1.95
F(9,69)	14.76	ARCH; F(1,67)	0.69
		Normality X^2 (2)	0.78
		Heterosced. F(16,52)	0.44
		Functional form F(26,45)	0.92
		Reset F(2,67)	3.43*

Diagnostic tests performed are those readily available in PC-GIVE, (see Hendry, 1989).

The resulting inflow of foreign exchange was channeled into housing market through the commercial banks. In addition, interest rates were rather low in 1987 in anticipation of a depression.

There were other reasons for the rise in households' demand for credit, such as increases in income resulting from better terms of trade. The increase in real bank lending was anyhow a symptom of serious disequilibrium in the housing market, caused partly by the rationing of rental housing and tax incentives for owner-occupied housing, such as the deductibility of interest payments on housing loans.

To account for a possible regime shift after the liberalization, which seems to be the main trigger for the rapid rise in asset prices after autumn 1986, we introduced a step dummy after 1986/3, money market liberalization, which however was not significant. This may not be surprising, since adaptation to new kinds of financial markets is not a once-for-all regime shift, rather an ongoing process that will take several years to accomplish. The explosion of asset prices in the first quarter of 1988, was accounted for by a single observation dummy D88/1. For housing prices there was a smaller outlier in 1986/2. The stock index had an outlier in 1985/3 and 1988/2. The recursive OLS estimations showed also signs of structural change in early 1980 s.

The major impact of real lending on real house prices is accomplished already within three quarters, but the variables are not cointegrated. The possible endogeneity of real bank lending as part of the financial allocation problem remains a bit of a mystery, since we did not get any robust results for the significance of interest rates.

6. Conclusions

We conclude that both asset price series have a unit root and are therefore integrated processes I(1). There was also a feeling that nominal house prices might even be I(2), but this was interpreted to reflect merely the exponential house price increases after the money market liberalization. The explosion in asset prices took place from autumn 1987 up until spring 1989.

The unit root property in the asset prices follows from the fact that weakly efficient markets reflect the stochastic trend in economic fundamentals. However, asset prices are not purely random walk with white noise errors, but rather unit root processes with stationary ARMA errors. The stochastic error part of these asset price processes is predictable from other variables, such as the evolution of other asset series and pricing errors seen in the error correction term. Therefore, this stochastic error contains the short run discrepancy from the market equilibrium.

In addition, some evidence was presented that the asset prices are cointegrated, which means that there exists a stationary linear combination of them, i.e. $ECT = (y_t - \theta x_t) \sim I(0)$. In fact, strong evidence for the cointegration hypothesis was found with the ADFR tests and Johansen tests for real asset prices.

Based on cointegration theory, an ECM model was constructed for asset price changes and several additional explanatory variables were tested for their effects on the slow adjustment in asset prices. Nominal asset prices also reflect to some extent the important arbitrage relation between the real risk adjusted, tax-free returns of assets, which motivates the cointegration hypothesis. However, problems in modelling nominal house prices due to the persistence of their nonstationarity remain, even after first differencing. As mentioned, these problems had a great deal to do with the money market liberalization and explosion of asset prices that followed.

Surprisingly, there was also a clear seasonal variation in house prices, especially in the first and second quarters. It seems natural to suppose that housing prices adjust in the short run whereas quantities adjust only in the long run. Under cointegration, the stochastic trend (unit root component) should carry the information about the change in long run market equilibrium. Cointegration does not have to hold, for example between two single stock price series, but it should hold between broad asset markets, if they are efficient enough. Therefore, cointegration also enables one to predict asset prices with the error correction mechanism.

A few additional variables were used in the error correction models of real house prices, among which real bank lending and stock market volatility, lagged permanent income and liquidity constraints seemed to have some value in predicting house prices.

References

Bengs, C. (1989), Pääkaupunkiseudun asuntojen hinnat vuosina 1965–1988, VTT tiedotteita 1007, Espoo.

Blanchard, O. & Fischer, S. (1988), *Lectures in Macroeconomics*, The MIT Press

Blangiewicz, M. & Charemza, W.W. (1990), »Cointegration in Small Samples: Empirical Percentiles, Drifting Moments and Customized Testing,» *Oxford Bulletin of Economics and Statistics*, 52, 303–315.

Bollerslev, T., Engle, R.F. & Wooldridge, J.M. (1988), »A Capital Asset Pricing Model with Time-varying Covariances,» *Journal of Political Economy*, 96, 116–131.

Bossaerts, P. (1988), »Common Nonstationary Components of Asset Prices,» *Journal of Economics Dynamics and Control*, 12, 347–364.

Campbell, J.Y. & Shiller, R.J. (1987), »Cointegration and Tests of Present Value Models,» *Journal of Political Economy*, 95, 1062–1088.

Campbell, J.Y. & Shiller, R.J. (1988), »Interpreting Cointegrated Models,» *Journal of Economics Dynamics and Control*, 12, 505–522.

Dickey, D.A. & Fuller, W.A. (1979), »Distribution of the Estimators for Autoregressive Time Series With a Unit Root,» *Journal of the American Statistical Association*, 74, 427–431.

Dickey, D. & Fuller, W. (1981), »Likelihood Ratio Statistics for Autoregressive Time Series with Unit Root,» *Econometrica*, vol. 49, p. 1057–1072.

Diebold, F.X. & Nerlove, M. (1988), »Unit Roots in Economic Time Series: A Selective Survey,» Federal Reserve Board,Finance and Economics Discussion Series, 49, in Fomby, T.B. & Rhodes, G.F. (eds.): *Advances in Econometrics: Co-Integration, Spurious Regressions, and Unit Roots*, JAI Press, Greenwich, Connecticut.

Engle, R.R. (1982), »Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of UK Inflation,» *Econometrica*, 50, 987–1007.

Engle, R.F. (1988), »Multivariate ARCH with Factor Structures — Cointegration in Variance,» Manuscript

Engle, R.R. & Granger, C.W.J. (1987), »Co-Integration and Error Correction: Repsentation, Estimation and Testing, *Econometrica*, 55, 251–276.

Engle, R.F., Lilien, D.M. & Robins, R.P. (1987), »Estimating Time Varying Risk Premia in the Term Structure: The ARCH-M model,» *Econometrica*, 55, 391–407.

Engle, R.F. & Yco, B.S. (1987), »Forecasting and Testing in Co-integrated Systems,» *Journal of Econometrics*, 35. 143–159.

Elton, E.J. & Gruber, M.J. (1987), *Modern Portfolio Theory and Investment Analysis*, 3. edition, John Wiley & Sons

Fuller, W.A. (1976), *Introduction to Statistical Time Series*, John Wiley, New York.

Granger, C.W.J. (1986), »Developments in the Study of Cointegrated Economic Variables,» *Oxford Bulletin of Economics and Statistics*, 48, 213–228.

Granger, C.W.J. (1988), »Causality, Cointegration and Control,» *Journal of Economics Dynamics and Control*, 12, 551–559.

Granger, C.W.J. (1988), »Some Recent Developments in a Concept of Causality,» *Journal of Econometrics*, 39, 199–211.

Hall, S.G. (1986), »An Application of the Granger and Engle Two Step Estimation Procedure to United Kingdom Aggregate Wage Rates,» *Oxford Bulletin of Economics and Statistics*, 48, 229–241.

Hendry, D.F. (1989), *An Interactive Econometric Modelling System, Version 6.0/6.01*, University of Oxford.

Hendry, D.F. & Neale, A.J. (1990), »A Monte Carlo Study of the Effects of Structural Breaks on Tests for Unit Roots,» Manuscript, Nuffield College, Oxford.

Johansen, S. (1988), »Statistical Analysis of Cointegration Vectors,» *Journal of Economic Dynamics and Control*, 12 p. 231–254.

Johansen, S. (1989), »Estimation and Hypothesis Testing of Cointegrated Vectors in Gaussian Vector Autoregressive Models,» Institute of Mathematical Statistics, University of Copenhagen, Preprint 1989 No. 3.

Lucas, R.E. Jr (1978), »Asset Prices in an Exchange Economy,» *Econometrica*, 46, 1429–1445.

Michener, R.W. (1982), Variance Bounds in a Simple Model of Asset Pricing,» *Journal of Political Economy*, 90 p. 166–175.

Nelson, C.R. and Plosser, C.I. (1982), »Trends and Random Walks in Macroeconomic Time Series — Some Evidence and Implications,» *Journal of Monetary Economics*, 10, 139–162.

Pere, P.J. (1990), »Integroituneet ja yhteisintegroituneet aikasarjat — Katsaus ominaisuuksiin ja testaukseen,» Elinkeinoelämän tutkimuslaitos, C 58, Helsinki.

Perron, P. (1988), »Trends and Random Walks in Macroeconomic Time Series, Further Evidence from a New Approach,» *Journal of Economics Dynamics and Control*, 12, 297–332.

Rothschild, M. (1985), »Asset Pricing Theories,» Manuscript La Jolla, UCSD, Department of Economics

Schwert, W.G. (1989), »Tests for Unit Roots: A Monte Carlo Investigation,» *Journal of Business and Economic Statistics*, 7, 147–159.

Suoniemi, I. (1990), »Uusien ja vanhojen asuntojen hinnat Suomessa, Hintasarjojen trendi- ja kausaalisuusominaisuudet,» TTT -tutkimusselosteita No. 103.

Takala, K. & Tuomala, M. (1990), »Housing Investment in Finland,» *Finnish Economic Papers*, No 2, p. 41–53.

Tarkka, J., Männistö, H-L. & Willman, A. (1989), »Consumption and Investment in the BOF4 Quarterly Model of the Finnish Economy,» Bank of Finland, Discussion Papers 9/89.

Appendix

Cointegration data: quarterly data 1970/1 – 1990/2 from the Bank of Finland model (BOF4) data

Variables

PHM	= Housing prices, whole country 85 = 100
PIH	= Construction price index, 80 = 100
LBP	= Bank lending to public, FIM million
R8870	= Effective dividend of the stock index
YD	= Household disposable income, FIM million
MTAX	= Marginal tax rate estimate for households
UNI	= Unitas price index for stocks, 75 = 100
RS	= Market rate of interest, %
RLB	= Bank lending rate to public, %
RB	= Government tax-free bond yield

Constructed variables

LIQCON	= (RS-RBL)*(LBP/YD); describes liquidity constraints
MRT-LBY	= (LBP/YD)*MTAX*RLB; relative tax advantage
MR-LBP	= MTAX*RBL*LBP; aggregate tax advantage
ECM	= Error correction term
D88/1, D86/2, D85/3, D88/2	= impulse dummies for single outliers
MM-LIB	= Money market liberalization step -dummy, value = 1 after 86/3
CV	= Total private consumption, FIM million (proxy for permanent income)
YD-INNO	= Income innovation term from RLS -estimation of disposable income or AR(1) plus constant and trend
UNI-ARCH	= Predictor of ARCH -variance from AR(5) -stock index difference equation

If not otherwise mentioned prefix L defers to logaritm, E for expectation, D for difference and R for real (deflated with appropriate price index). As in PC-GIVE notation 1 in the end of the variable refers to one period lag.