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House prices and long-term equilibrium in the regulated market of the Netherlands

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ABSTRACT

This paper establishes a simple affordability model that implicitly incorporates the major Dutch market features to elucidate long-run house prices under a regulatory environment. The results reveal a long-run relationship for house prices under strict regulations. The association among house prices, income, interest rates, and inflation is verified using an aggregated dataset. In the long-run, incomes and interest rates function as the two prime forces driving price dynamics, whereas the role of inflation is limited.

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House prices; regulations; the Netherlands; long-run equilibrium

1. Introduction

After the sharp rise of the 1990s, house prices in the Netherlands started to grow moderately and reached a peak in 2008 followed by a significant decline. Unlike most other European countries, house prices in the Netherlands experienced a longer period of drops. The market showed no signs of recovery until the last quarter of 2013. The unexpectedly divergent pattern draws attention back to the long run relationship of house prices before the crisis. Was the peak price caused by the deviation from the equilibrium that immediately led to a 'spontaneous' market adjustment in the form of the crash? Or was it a result of a bubble-generating process? This paper investigates the long-run house-price relations in the Netherlands from 1982 to 2008, which offers a great opportunity to delve into the long-run equilibrium in a regulated context. Ranked among the most highly regulated markets, the Netherlands has a predominant social housing rental sector, a strongly subsidized housing market, and a highly inelastic supply sector. These together contribute to the, somewhat special, price path in the Dutch market. The pre-crisis period 1982–2008 was selected because it experienced no external shocks, possibly providing insights into the connection with the recent global financial crisis.[1]

Long-run equilibria are widely used to investigate the extent to which house prices are over- or under-valued, further explaining real estate booms and bursts. This strand has focused on the interplay of demand and supply (see, e.g. Chen & Patel, 1998; Hott & Monnin,

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2008; Ambrose, Eichholtz & Lindenthal, 2013). General models like the asset pricing house-price model have been developed for markets like the US, where government intervention is limited (see, e.g. Poterba, 1984; Himmelberg, Mayer & Sinai, 2005; Brunnermeier & Julliard, 2008; Gallin, 2008; Campbell et al., 2009; Mols & Lindenthal, 2011). However, these models can run into problems when it turns to the regulated markets[2] as the underlying assumptions are not satisfied.

The price-to-income ratio and rent-to-price ratio are broadly used as indicators to measure the affordability in housing markets. However, the two ratios hardly found fertile soil in the Dutch market research, considering the regulatory environment. As argued by Himmelberg, Mayer, & Sinai (2005), differences in intrinsic market characteristics need to be considered when discussing market heterogeneity. Given the favorable tax deduction policy, the price-to-income ratio does not fully account for the fiscal benefit. The rent-to-price ratio, on the other hand, is established presuming a competitive market and does not fit in regulated markets (Ayuso & Restoy, 2006). The mismatch can be easily perceived from the Dutch market from the large discrepancy between rent increases and price growths which could have been further fueled by the rent regulations that imposes a ceiling for rent increases through the 'spill-over' effect.

The concept of the interest-to-income ratio (IIR) seems particularly well suited to the Netherlands. It accounts for affordability under favorable fiscal policy and potentially captures price dynamics more precisely (Boelhouwer *et al.*, 2001, 2004). Although the IIR-based model implicitly embodies an equilibrium relation, whether evidence of an equilibrium exists remains unclear. We examine deeper the long-run equilibrium of house prices by extending the IIR into a flexible long-run relationship. We attempt to verify the relationship empirically using aggregated national data from 1982 to 2008. This period excludes the regime-switching phases due to the crisis to avoid further complexity. We use an error correction model to explore how the detected long-run relation performs in a dynamic housing market.

This paper builds upon previous studies of the long-run equilibrium relation of IIR. Although we do not specify the definition and the measurement of regulation as it is not the aim of this study, the regulatory environment has been considered implicitly in the setup of the long run equilibrium. The regulatory aspects of the Dutch market are also used to interpret the empirical findings. To our knowledge, we are among the first to ascertain a flexible house-price equilibrium centering on affordability in the Dutch market. The rest of this paper is organized as follows. Section 2 reviews house-price equilibria and describes the Dutch market. Section 3 explains the house-price model based on the IIR, followed by demonstrating an extended model that captures a more flexible long-run relation. The methodology is presented in Section 4. Section 5 provides the data. The modeling results and discussion are presented in Section 6. The final section concludes.

2. Models of long-run equilibrium

2.1. Competitive markets

House-prices in the long-run have been extensively modeled and applied to competitive markets. Theoretical literatures have mainly centered on the neoclassical approach that emphasizes supply and demand. In an 'ideal' world with perfect elasticity of supply, long-run equilibrium is solely determined by supply-side factors: typically, construction costs and land prices (see, e.g. Madsen, 2012). The underlying assumption of complete elasticity is, however, barely satisfactory in the real world. Ignoring the influence of demand can lead to potential problems in a short- and medium-term scenario. In that regard, most studies consider not only the supply but demand-associated fundamentals to investigate the equilibrium.

Rent-to-price and price-to-income ratios are widely used as the proxy of equilibrium (Buckley & Ermisch, 1982; Muth, 1988; Capozza & Seguin, 1996; Cho, 1996; Gallin, 2008; Davis, Lehnert & Martin, 2008). Like the dividend-price ratio initiated for the stock market, the rent-to-price ratio is grounded in asset-pricing theory. In a similar vein, households compare the long-term return of each tenure choice — renting or owning — before reaching a decision (Poterba, 1984). Any opportunity for arbitrage would be eventually squeezed out, bringing house prices back to equilibrium. Such a mechanism implies that markets will adjust quickly and is more appropriate to competitive markets.

Compared to the rent-to-price ratio, less consensus has been reached in explaining the mechanism of the price-to-income ratio. There are, however, at least two ways to imply the connection. According to the law of consumption, the excess housing demand stimulated from the rise of incomes would in turn impose upward pressure on house prices. The second channel associates house prices to lending criteria (see, e.g. Tsatsaronis & Zhu, 2004; Gerlach & Peng, 2005; Iacoviello, 2005). Financial intermediaries regard the price-to-income ratio as an indicator to evaluate household's ability of repaying debts. Thereby, they require a proportional increase in household income for a larger size mortgage (Hulchanski, 1995). There is a large body of literature on this long-run relationship (see, e.g. Abraham & Hendershott, 1996; Malpezzi, 1999; Capozza *et al.*, 2002; Meen, 2002; Black *et al.*, 2006). For example, using a sample of American cities from 1978 to 1992, Abraham & Hendershott (1996) set up a linear equilibrium model for real house prices and income. They reported a price elasticity for income of 0.50. Though broadly used, this ratio lacks in an adequate theoretical basis when compared to the rent-to-price ratio.

The demerits of these two ratios have been discussed by Himmelberg *et al.* (2005). They argue that neither ratio would appropriately reflect the long-run association partially due to the changes in actual cost of housing that could arise in regulated markets. On the other hand, many studies tried to provide empirical evidence through establishing long-run equilibria for house prices by positing a relationship between house prices and fundamentals (Abraham & Hendershott, 1996; Capozza *et al.*, 2002), accounting for impacts of different aspects such as household wealth and bank lending behavior (Drake, 1993; Gerlach & Peng, 2005; Gallin, 2006; Zhou, 2010). Some researchers have offered evidence of long-run relationships based on econometric tests, though the results are inconclusive. Abelson *et al.* (2005), for instance, performed a multivariable cointegration test and found a positive association among the real house price, income, and the consumer price index. Contrarily, Gallin (2006) failed to demonstrate a long-run relationship between house price and income based on the result of a bootstrap approach using US data, and questioned the validity of the conventional price-to-income ratio.

2.2. The Dutch market in a regulated context

Like most European housing markets, the Dutch market is highly regulated and differs from the competitive markets substantially as government interventions, including supply constraints and housing subsidies, have largely reshaped the owner-occupied market. In fact, the government has intervened in almost every respect (van der Klaauw & Kock, 1999). Policies and legislations were introduced to enhance ownership and to improve housing affordability. This has led to a housing stock comprised of a highly stimulated owner-occupied sector, about 60% of the total stock in 2012, and a small private rental sector of 8%, not to mention that even part of the private rental sector is also regulated.

Besides providing a direct subsidy to low-income tenants, the Dutch government largely subsidizes owner-occupancy through tax policy. Ranked among the most generous fiscal rules, the policy allows households to deduct the mortgage interest from their taxable income for maximum 30 years. Accelerated by the financial liberalization that brought about a diversity of funding choices in the 1990s, this policy has effectively stimulated the housing demand.

Given the severe scarcity of land, the Netherlands has adopted stringent spatial planning laws and zoning rules. These constraints, together with the requirement of additional legal permits, prolong the construction process, adding difficulty to expand the housing stock. As expected, housing supply becomes highly inelastic. In fact, this sector was described as fixed in size, given the small amount of newly built housing compared to existing stock (Boelhouwer *et al.*, 2004).

These constraints have engendered a special pattern of house-price dynamics. Under rent controls, excess demand would be stimulated for the rental units as it becomes relatively cheaper. When the excess cannot be met with sufficient rental units, it will be transferred to the owner-occupancy sector, pushing up the house prices (Gould & Henry, 1967). This implies that rent control is equivalent to subsidizing households by increasing households wealth[3], shifting demand to the owner-occupied market (Koning & Ridder, 1997). Scholars have, however, augured the ambiguous effects from rent controls as the supply would adjust accordingly and that the ultimate impact on house prices is determined by the extents to which both the demand and supply curve shift (see, e.g. Gould & Henry, 1967; Wang, 2011). The highly inelastic supply sector would also impose additional positive pressure on house prices, stimulating a upswing market, and thereby adding additional risk to the system, which in turn makes the financial system more fragile. In this way, constraining supply may exacerbate the real estate cycle (Malpezzi & Wachter, 2005).

Limited literature has studied the long-term house prices in the Netherlands. Two pioneering studies were carried out by Boelhouwer *et al.* (2001, 2004) and Kranendonk *et al.* (2005). The latter empirically discussed the determinants of house prices and reported a long-run relationship among house prices, income, interest rates, financial assets, and housing supply. This empirical study was re-estimated with an extended sample by Kranendonk and Verbruggen (2008), who reported evidence of the long-run equilibrium. Instead of examining the empirical effects of each fundamental, Boelhouwer *et al.* (2001, 2004) proposed an interest-to-income ratio to represent the equilibrium, which was embedded within an error correction model to reflect the market dynamics. They argued that the model particularly fits the Netherlands as the 'borrowing limit' which is dependent of the income levels actually determines the house prices. Viewing this long-run ratio as a good reflection of the Dutch market, we develop a simple flexible house price model based on the ratio.

Other studies, including investigations by the OECD (2004) and Hofman, Nadal-De Simone, & Walsh (2005), started from an empirical point. In addition, Francke, Vujic, & Vos (2009) compared existing models and favored the unobserved component error-correcting model with random walk. Briefly, the existing literature on Dutch house prices emphasizes the empirical work.

3. A simple flexible model

In this section, we propose a simple flexible long-run relationship of house prices based on a demand-oriented model from the perspective of household behavior.

3.1. The long-run relationship of house prices

The long-run relationship of house prices is proposed as

$$\ln P_t = b_0 + b_1 \ln Y_t + b_2 \ln i_t + b_3 \pi_t, \tag{1}$$

where b_i's are the parameters, π_t denotes the inflation rate, P_t is the average house price, Y_t represents the household income, i_t is the mortgage interest rate.

Equation (1), indicates that house prices over time are determined by the income level, interest rate, and inflation. Appendix 1 illustrates the derivation of such a flexible relationship from a simple demand-oriented model. The model considers a representative household who consumes housing and a composite good to maximize his utility subject to a budget constraint. This household gains a separable utility through consuming both housing and the composite good, with a constant elasticity of substitution of the intertemporal consumption of the two goods.

The household also faces a periodic budget constraint as spending on consumption has to be balanced with the repayment on the mortgage loan through income that takes account of the benefits generated from tax relief policy. We also assume that the amount of mortgage repayment (both the amortized amount and interest) on housing in each period is a fixed fraction of the total loan, which is further expressed by a nonlinear function of tax relief factor and interest rates. We acknowledge that this strong assumption ignores the repayment schemes originated in different types of mortgage contracts (see Chambers *et al.*, 2009). This simplification, however, does reflect the aggregated mortgage market to a certain extent. First, it resembles well the repayment scheme of the interest-only mortgage, wherein repayment tends to be postponed until the loan matures. The fixed repayment assumption also applies to the annuity mortgage. Given that most other mortgage types, except for graduated payment mortgages, have an initial payment higher than that of the interest-only mortgage, these mortgage contracts require more capital to be built up in the early stages across the mortgage schemes; this may in turn add extra burden to households who face a credit restraint. In that regard, with our assumption it makes sense to consider the most favorable case, especially for the households who are driven to become a homeowner. We recognize that this may bias upwards house prices in our model.

To simplify the setup, we additionally assume this household consumes only one unit of housing and presents no moving behavior. Solving the maximizing problem leads to the flexible house-price relationship expressed by the interest rate and expenditure on the composite good. The latter is further decomposed as a fraction of household's weighted income. By approximating the income growth rate, the flexible long-run house price is expressed by income, interest rates, and inflation. The derivation can be found in Appendix 1.

3.2. Linkage to an existing Dutch model: interest/income ratio

The long-run house price in Equation (1), is closely connected to the house-price model by Boelhouwer *et al.* (2001, 2004) (hereafter, the OTB model). They first introduced the interest-to-income ratio (IIR) to reflect the long-run equilibrium, which was subsequently extended to incorporate short-term effects. The IIR is a reasonable proxy of the long-run equilibrium in light of the favorable tax policy that inspires households to purchase houses, the borrowing limits imposed by financial institutions, and the high loan-to-value ratio observed in this market (de Vries & Boelhouwer, 2009).

As an intrinsic affordability model that highlights the role of households, the OTB model suggests that income and the interest rate comprise the essential determinants of house prices in the long run, which implies that house prices are effectively affected by households' behavior rather than supply factors. Tax benefit has been additionally incorporated into the IIR by introducing a tax-advantage multiplier. The long-run ratio of interest-to-income (c) presented in the OTB model can be expressed as:

$$c = \frac{P_t i_t (1 - F_t)}{Y_t}, \tag{2}$$

where P_t is the average house price, Y_t represents the household income, i_t is the interest rate, and F_t is the tax relief factor. We use the same constant tax relief factor 0.405 as Boelhouwer *et al.* (2004). This is, however, a strong assumption as the marginal tax rates have varied during 1982–2008. Due to data limitation, we are not able to calculate the actual tax margins as time series . To gain extra hints from data available, we tried data from the Dutch Housing Survey (WoON) 2002, 2006 and 2009 to recalculate the weighted marginal tax rates for home owners under 64 years who carry mortgage loan. The tax relief factor stayed within the narrow range of [0.40, 0.42]. If data permits, the accuracy of the results should be improved given a more accurate measure of tax relief factor.

The proposed interest-to-income ratio (IIR) fairly reflects the interest-only[4] mortgage scheme. Consider a household that borrows to finance their house

$$L + D = P, \tag{3}$$

where L is the mortgage loan while D and P denote the down payment of the asset and house price, respectively. A down payment ratio $(1 - \theta)$ is taken as a fraction of the house price. As financial institutions expect the household to pay a monthly interest that is assumed as a portion of income, it makes sense that a fixed fraction of the household's income is reserved to pay the monthly mortgage interest. Given the fiscal benefits, the income budget is expanded through the tax deduction. A household's budget therefore becomes

$$i_t L_t = \varphi Y_t + F i_t L_t, \tag{4}$$

where F is the tax relief factor (0.405), i denotes the interest rate, φ is the fixed fraction of income spent on the periodic mortgage interest payment, Y represents household income, and the subscript t indicates the time period.

Substituting L_t by $P_t \theta$ gives the interest-to-income ratio

$$\frac{\varphi}{\theta} = \frac{P_t i_t (1 - F)}{Y_t}, \tag{5}$$

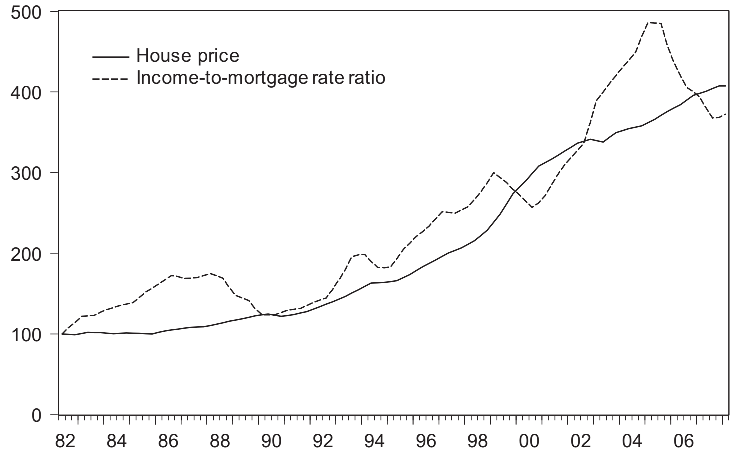

Figure 1. Nominal house price and the income-to-interest rate ratio (1982Q2 = 100).

which is equivalent to IIR, except that the ratio becomes $\frac{\varphi}{\theta}$. Recall that φ is the fraction of income spent on repayment, and θ is the loan in percentage of house price. This term indicates that a higher income is required to compensate for the increased default risk if a household attempts to contract for a larger loan.

Data observed from the real market corroborate the pattern suggested by the IIR ratio. Figure 1 depicts the house prices and the inverse of IIR from 1982 to 2008 as the inverse (i.e. income-to-mortgage) provides clearer trend compared to house prices. These two series move closely, suggesting a similar trend throughout the period, despite that deviations occur in the short run.

However, the long-run equilibrium in the form of IIR implicitly imposes several strong assumptions. Exogenous quantitative relations have been set up *ex ante* between prices, incomes, and interest rates. This can be easily shown by taking the logarithms of both sides of Equation (2):

$$\log(P_t) = c_1 + \log(Y_t) - \log(i_t), \tag{6}$$

where c_1 is the constant $\log\left(\frac{c}{1-F}\right)$. This form states a long-run unit price elasticity of both income and interest rate, which lacks a theoretical explanation.

As illustrated earlier, the long-run equilibrium in the form of IIR is a restricted form of the flexible long-run relationship in Equation (1).

3.3. The role of inflation

Apart from the role inflation played in affecting income growth (see Appendix 1), three additional explanations of the role of inflation are provided. First, using a nominal measure[5] for all of the variables indicates that inflation may still have an impact in the long run as the variables are likely to be affected differently by inflation. In this regard, inflation serves as a bridge to associate the macroeconomic environment with the households' behavior. This can be seen, for example, from the phenomenon of money illusion: households may be misleadingly guided to favor renting over ownership simply because they do not take inflation into account.

The user-cost view suggests that inflation could have an impact on real prices because it enters the fundamental user cost through real price growth (Poterba, 1984; Meen, 1990). User cost (ψ_t) is composed of the housing maintenance cost δ, interest cost $(1 - F)^* i_t$, and the nominal house-price appreciation, which is the sum of real house-price growth (π_t^{Pr}) and inflation (π_t). The effect of inflation can be easily shown by taking the partial derivative of the user cost with respect to inflation:

$$\frac{\partial \psi_t}{\partial \pi_t} = (1 - F)\frac{\mathrm{d}i_t}{\mathrm{d}\pi_t} - \frac{\mathrm{d}(\pi_t^{Pr} + \pi_t)}{\mathrm{d}\pi_t}. \tag{7}$$

Equation (7) indicates that inflation would impose a upward pressure on house prices as the cost of owner-occupancy is further reduced due to the presence of a positive inflation. Although the stimulated demand may be partially mitigated by increasing the housing supply, it cannot be completely eliminated because of the non-perfect elasticity of supply. We additionally performed a long-run flexible relation in the real terms that are generated by deflating the nominal variables with the consumer price index in the base year of 2006, to compare the effects of inflation.

The third reason comes from the empirical result by Kranendonk *et al.* (2005). Although inflation has not been considered in their long-run relationship, the error correction modeling results present significantly different impacts from changes in nominal interest rate and inflation. The estimated coefficient of change in the nominal mortgage rate is about four times as large as that of inflation. That result is not strictly consistent to the indications of the long-run relationship. Thus, it may be reasonable to adapt the long-run relationship by adding the impact of inflation.

In summary, we have extended the restricted interest-to-income ratio to a flexible relationship that gives additional emphasis on the impact of inflation. The empirical part mainly focuses on the long-run relation of Equation (1). The long-run flexible relationship derived from the simple model well fits the regulated Dutch market, as it takes account of the regulatory effects. This demand-oriented setting closely reflects the highly inelastic housing supply in the Netherlands. From the perspective of households, the budget constraints indicate how they benefit from the favorable tax policy by selecting an appropriate mortgage scheme. Despite the strong assumption that we have imposed on the repayment, the model still could reasonably represent the regulated market of the Netherlands. Loosely worded, we do not aim at providing precise measurement of regulatory environment, but implicitly incorporate such features into the model setting process.

4. Method

We adopt the cointegration analysis for the empirical part. The concepts of (non)stationary series, unit root, and integration are introduced first. Stationarity (or nonstationarity) is determined by the probability distribution of the series.[6] On contrast to a stationary series that temporarily fluctuates but eventually tends to return to its mean, nonstationary series will accumulate any minor changes, suggesting an indecisive pattern in trend. Unit root is equivalent way to measure the nonstationary process using the root of the characteristic equation. The third is the integration order, which describes how far a nonstationary series deviates from a stationary process (Granger, 1981).

Although single nonstationary series will not revert to a certain level, multiple series can share a similar pattern which is preferably indicated by economic theory. Cointegration describes such an intrinsic natural association among different variables in the long run. In this study, we attempt to find the cointegration relationship among house prices, interest rates, incomes, and inflation.

We apply both the single equation method proposed by Engle & Granger (1987) and a system dynamic method proposed by Johansen (1988, 1991) to identify cointegration. These two methods differ in treating the variables of interest: the Engle–Granger method gives focus to a single cointegrated equation, and the VECM method views all the variables as an endogenous system. They are both presented to testify to the existence of cointegration in this study.

The Engle–Granger method considers the long-run relation as a single equation which is used in the second stage to combine the short-term dynamics. This allows us to reach an error correction model (ECM) that stresses the effect of the long-run attractor. Generally, the times series of interest Z_t can be represented in error correction form if the cointegration relation exists among the variables:

$$A(L)\Delta Z_t = \lambda e_{t-1} + B(L)\Delta X_t + \mu_t,\ Z_t = a_0 + \boldsymbol{a}'\boldsymbol{X}_t + \boldsymbol{e}_t,$$

where $A(L)$ and $B(L)$ are the lag operators, $\boldsymbol{X}_t$ is the $k \times 1$ vector of the k explanatory series, $\boldsymbol{a}'$ is the corresponding parameter vector, $\boldsymbol{e}_t$ is the vector of the error term, and λ and a_0 are non-zero constants. In the first step, the series with the same integration order can be applied directly using ordinary least squares, which provides consistent[7] estimates.

The Johansen test, in contrast, regards the series as a system. The Johansen procedure proposes to examine cointegration for multiple-series systems based on reduced rank regressions. A simple vector error correction model can be formed from the vector autoregressive model (VAR) form

$$\boldsymbol{Y}_t = \boldsymbol{\Phi}_1\boldsymbol{Y}_{t-1} + \boldsymbol{\Phi}_2\boldsymbol{Y}_{t-2} + \cdots + \boldsymbol{\Phi}_p\boldsymbol{Y}_{t-p} + \boldsymbol{b}_t + \boldsymbol{\varepsilon}_t$$

where $\boldsymbol{Y}_t$ is the $m \times 1$ vector of the system series, $\boldsymbol{\Phi}_i$ $(i = 1, 2, \cdots, p)$ represents the coefficient matrix, $\boldsymbol{b}_t$ is a constant vector, and $\boldsymbol{\varepsilon}_t$ is the error term.

When cointegration exists among the system $\boldsymbol{Y}_t$, the above VAR presentation can be arranged as a simple vector error correction model (VECM)

$$\Delta\boldsymbol{Y}_t = \boldsymbol{\Pi}\boldsymbol{Y}_{t-1} + \sum_{i=1}^{p-1}\boldsymbol{\Phi}_i^*\Delta\boldsymbol{Y}_{t-i} + \boldsymbol{b}_t + \boldsymbol{\varepsilon}_t$$

where $\boldsymbol{\Pi}$ is expressed by $-(\mathbf{I} - \boldsymbol{\Phi}_1 - \boldsymbol{\Phi}_2 - \cdots \boldsymbol{\Phi}_p)$, $\boldsymbol{\Phi}_\mathbf{i}^*$ equals to $-\sum_{n=i+1}^{p}\boldsymbol{\Phi}_\mathbf{n}$, with $n = 1, 2, \cdots$ (p–1). The matrix $\boldsymbol{\Pi}$ determines the cointegration properties of the system. When the rank of matrix $\boldsymbol{\Pi}$ —assume it is denoted by r —is between 0 and m, then the matrix $\boldsymbol{\Pi}$ can be easily decomposed into

$$\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$$

where the columns of $\boldsymbol{\beta}$ shows the r cointegrating vectors. In this system, the long-term equilibrium can be achieved by setting $\Delta \boldsymbol{Y}_t = \mathbf{0}$.

Both methods have prerequisites to ensure the reliability of the modeling results, including that the series should either be stationary or share a same order of integration[8]. It is still possible to analyze nonstationary series with the definition of integration (I) provided by Granger (1981). According to that definition, a nonstationary series could possibly be converted to a stationary series after several times' differencing (I(d)). Granger (1981) claims that the spurious estimation induced from nonstationary series can be excluded if a constant vector can be found that ensures a certain combination of nonstationary series to be stationary. Nonstationary series are said to be cointegrated when the constant vector exists.

In practice, nonstationarity has been examined by conducting unit root tests based on a standard Dickey–Fuller (DF) test. We mainly use an improved unit root tests: the augmented Dickey–Fuller (ADF) test derived from a parametric approach. The ADF tests provides three forms:

$$y_t = \rho y_{t-1} + \alpha + \beta t + \sum_{1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t,$$

$$y_t = \rho y_{t-1} + \alpha + \sum_{1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t, \text{ and}$$

$$y_t = \rho y_{t-1} + \sum_{1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t.$$

However, no consensus has been reached on how to choose the appropriate test form. This study goes on to determine the unit root mainly on the basis of ADF results. Where results were close to the critical values, the Ng-Perron test[9] (NP test) was applied to reach a conclusion.

5. Price development and the data

We use quarterly data from 1982Q2 to 2008Q1. This sample period was selected for the reason that house prices during this particular period showed a general increase. This upswing provides a relatively smoothing context to examine how house prices develop without external shocks and to establish the corresponding equilibrium. This period, however, experienced changes in policies that are associated with lending restrictions: for example, the second earner income was considered since 1993 and the tax deduction was set to a maximum 30 years. A detailed chronological review of such changes were discussed by Elsinga, Priemus, & Boelhouwer (2016). These changes can potentially affect the house prices through the lending channel; it is, however, extremely difficult to identify the influence of single policy change such as the impact of stricter lending criteria due to the interplay between the mortgage interest rates and house prices (see, e.g. Rappoport, 2016; Cloyne

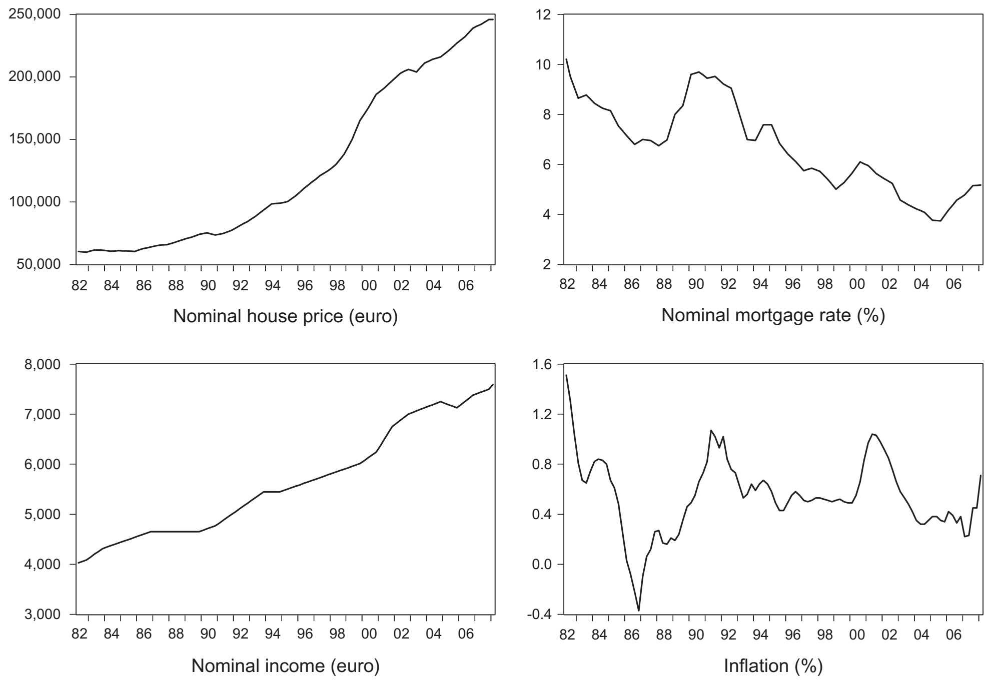

Figure 2. Series of interest in nominal terms (1982Q2–2008Q1).

et al., 2017). We therefore did not tract the exact effect of a certain policy in this study, but used a flexible long-term setting that potentially allows for changes in policies in the form of a reduced equation.

Average house-price was obtained from the Netherlands' Association of Realtors (NVM) which offers the average price of existing housing units, excluding newly built dwellings. Information on the mortgage market comes from the National Bank (DNB). The inflation[10] figures are available from the Statistics Netherlands (CBS). The Netherlands Bureau for Economic Policy Analysis (CPB) publishes household incomes annually. This average gross income is generated from about 80% of gross average (household) income.[11] We use the linear match method to interpolate the income series into quarterly base. Data used in this study has several disadvantages such as using information on an average level means that heterogeneity is ignored.

As shown in Figure 2, both house price and income series share a rising trend, but the former grows more rapidly with house prices tripling in the past twenty years. This divergence suggests that income alone cannot sufficiently explain the booming house prices. The next few years after the second oil shock presents a modest recovery in house prices until 1990 when a dramatic growth took off. Along with the tremendous increase in price, the interest rate plummeted; for instance, the nominal interest rate has reduced by over two-thirds at its trough. The large variation of interest rate ranged from a ceiling of 10.2% to a minimum 3.7%, which may have spurred the housing market. The series of inflation data indicates a significant decline during the first five years and a roughly stable fluctuation for the remaining period. Table 1 and 2 summarize the data descriptions.

Table 1. Descriptive statistics of nominal series (1982Q2–2008Q1).

Variable	$\ln P_t$	$\ln Y_t$	$\ln i_t$	π_t
Mean	11.62	8.62	1.86	0.54
Median	11.51	8.61	1.92	0.53
Maximum	12.41	8.94	2.32	1.51
Minimum	11.00	8.30	1.32	−0.37
Std. Dev.	0.50	0.19	0.27	0.30

Table 2. Descriptive statistics of series in real terms (1982Q2–2008Q1).

Variable	$\ln P_t^r$	$\ln Y_t^r$	$\ln i_t^r$	π_t
Mean	11.37	8.37	1.77	0.54
Median	11.26	8.36	1.85	0.53
Maximum	12.44	8.96	2.20	1.51
Minimum	10.48	7.78	1.21	−0.37
Std. Dev.	0.67	0.35	0.28	0.30

Notes: The superscript r denotes the real term. Real house price and real income are generated by deflating the CPI in the base year of 2006. This is the mostly used base year by CBS. The real interest rate is calculated by $(1 + i)/(1 + \pi) - 1$.

Table 3. ADF test based on the observed deterministic components.

Variable	$\ln P_t$	$\ln Y_t$	$\ln i_t$	π_t
Form	Drift, trend	Drift, trend	Drift, trend	Drift, trend
Lag length	1	1	1	4
p-value	0.63	0.09[a]	0.16	0.07[a]
Unit root	Yes	Yes	Yes	Yes
I(d)	I(1)***	I(1)**	I(1)***	I(1)***

Notes: The lag length is selected based on SIC. I(d) indicates the integration order. Asterisks '***', '**', '*' indicate a significance level of 1, 5, and 10%, respectively.
[a]The ADF test results of $\ln(Y_t)$ and inflation indicate no unit root at 9% and 7% (close to the critical values), which cannot be approved by Ng-Perron tests at 10%.

6. Results and discussion

6.1. Unit root and integration

The ADF test results in Table 3 cannot reject the existence of a unit root for house prices and mortgage rates, indicating nonstationarity. However, the results of incomes and inflation reported no evidence of unit root at a significance level close to 10%. This gives us the reason to further conduct the Ng-Perron tests, which confirmed the existence of unit root of income and inflation. Repeating the tests for the first order difference, we found evidence that all the four series are stationary at least with a 5% significance level. Therefore, we conclude that all the series are nonstationary, but integrated at order one.

6.2. Long-run relationship

6.2.1. Long-run equilibrium as in the single equation

Given the common integrating order shared by the series, the long run equilibrium is further investigated.

Two outputs are provided based on the Engle–Granger procedure: the long-run relationship in Table 4 and the associated error correction model generated from the Engle–Granger

Table 4. Output of the flexible long-run relationship (dependent variable = ln P_t).

Variable	Coefficient
C	−7.57***
ln Y_t	2.28***
ln i_t	−0.31***
π_t	0.18***

Notes: Asterisks show a different significance level, particularly '*', '**', and '***' indicate a significance level of 10%, 5%, 1%. The output also reports an adjusted R^2 of 0.98 and a Durbin–Watson statistic of 0.11.

Table 5. ECM output of EG (dependent variable = Δ ln P_t).

Variable	Coefficient
Correction	−0.02*
Δ ln Y_t	0.29**
Δ ln P_t(−1)	0.84***
$\bar{R}^2 = 0.63$	DW stat = 1.85

Notes: Asterisks '*', '**', and '***' indicate a significance level of 10, 5, and 1% respectively.

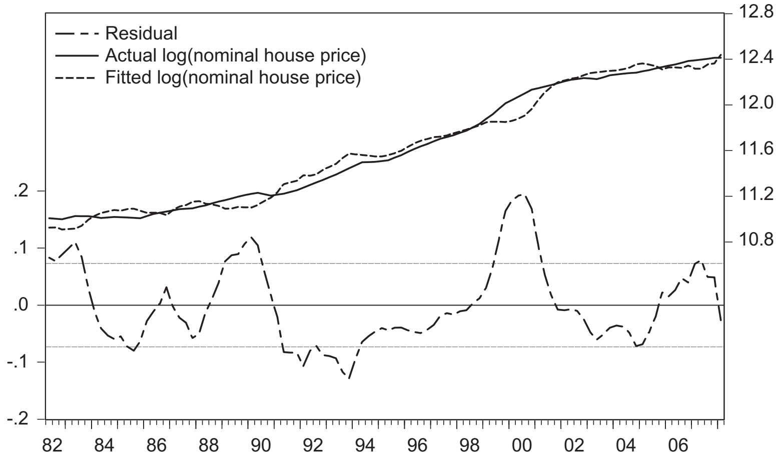

Figure 3. Residuals generated from the flexible long-run relationship.

procedure in Table 5. The residuals derived from this relationship (in Figure 3) suggest a reverting trend which was further confirmed by ADF test at the 1% significance level.

Figure 3 also shows the gap between the long-run house prices and the actual house prices. The two share a similar trend during the sample period. Despite the frequent deviations in the short-term, actual house prices tend to return to the equilibrium, which is most likely to be driven by the attractor of the long-term prices.

The associated ECM form was chosen based on a general-to-specific process in which a standard ECM with 2 lags was applied first. The results are reported in Table 5. The estimation reports an adjusted R^2 of 0.63 and an uncorrelated error term that has been confirmed by the Breusch-Godfrey LM test of two lags. The estimated long run relation significantly corrects the short term deviation by 2% per period. Figure 4 compares the house prices in logarithm estimated from the ECM with that of the actual house prices, indicating a fair fit.

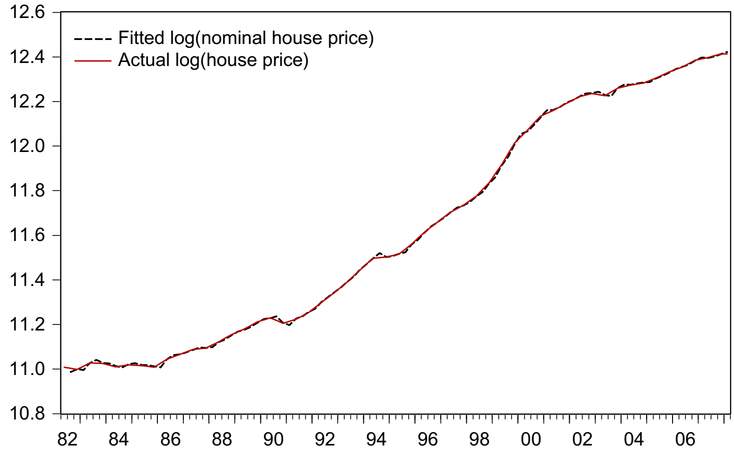

Figure 4. The model estimated house price and the actual house price in logarithm.

Table 6. Johansen cointegration test result (trace test and max-eigenvalue test).

Trace test			Max-eigenvalue test		
H_0:cointegration numbers	Eigenvalue	Probability	H_0:cointegration numbers	Eigenvalue	Probability
None *	0.29	0.00	None*	0.29	0.00
$r \le 1$	0.19	0.06	$r \le 1$*	0.19	0.03
$r \le 2$	0.06	0.60	$r \le 2$	0.06	0.53
$r \le 3$	0.00	0.67	$r \le 3$	0.00	0.67

Notes: Asterisk '*' indicates that the null hypothesis of maximum r cointegration relationships is rejected at the 5% level.

6.2.2. Long-run equilibrium in the system

We also tested the long-run relationship of house prices viewing the variables as an endogenous system in a VAR form. Different from the Engle–Granger method, the dynamic system generates the long-run relationship endogenously, which allows performing the error correction model in a higher dimensional space (vector error correction model). According to the Johansen test results in Table 6, both the trace test and max-eigenvalue test reject the null hypothesis of zero cointegration, but the trace test cannot reject the null of maximum one cointegration, suggesting the existence of a cointegration relation.[12] Based on the AIC and SC statistics at 5%, a lag order of two of the Johansen tests were implemented in a VAR system, which means the dynamic system could be captured in the vector error correction form with one lag.

We report the cointegration result of VEC (1) in Table 7. The VEC (1) model generates an adjusted R^2 of 0.66 slightly higher than the EG result. The error term in this system, however, insignificantly corrects house-price dynamics.

The cointegration relation generated by the EG method and that from VEC analysis suggest a similar pattern. We choose to focus on the former from onwards because the parameters of EG method are more straightforward in their economic meaning.

Strong empirical evidence of a long-run relationship amongln P_t, $\ln i_t$, π_t and $\ln Y_t$ is reported in Table 4. We found positive associations for household income and inflation in contrast to the impact of the interest rate. Particularly, a one-percent increase in income

Table 7. Cointegration derived from VEC (1) output (dependent variable = ln P_t).

Variable	Coefficient
C	−1.49**
ln Y_t	1.37***
ln i_t	−0.99***
π_t	0.47***

Notes: Asterisks '**', and '***' indicate a significance level of 5% and 1% respectively.

would stimulate more than 2% rise in house price. The magnitude of the price-to-income elasticity is within the vicinity estimated by international studies, ranging from 0.2 to over 8.0. Focusing on the metropolitan areas in the US, Capozza *et al.* (2002) found an elasticity of 0.45 using a panel sample from 1979 to 1995. Abraham & Hendershott (1996) also reported less than unit elasticity based on empirical results from US data. Abelson *et al.* (2005) obtained an income elasticity of 1.7 for the Australian housing market. A larger elasticity of income was found by Meen (2002), who reported 2.54 for the UK market for the period 1969 to 1990 and 2.71 for the US market from 1981 to 1998. A more recent study by Ott (2014) found a price elasticity of income of 1.91 for Euro area from 1970 to 2012.

Compared to literatures on the Dutch market, our price elasticity of income is slightly higher than the upper range of 2, albeit comparable. The empirical literature indicates no consensus on the impact of income in the Dutch market. The sequence of studies from CPB claimed an income elasticity around 1.5, and showed the impact of income has increased by 20 basis point when the sample period 1981–2003 was extended to 1980–2007. A smaller influence of income (around 1) has been reported by Francke, Vujic, & Vos (2009) with a sample period 1965–2009. The unit income elasticity has been implicitly presumed by de Vries & Boelhouwer (2009). The variation in existing studies indicates a sample-period dependent influence of income. Our result is more close to the work by OECD (2004), which reported 1.94. The relatively larger income elasticity can be partially explained by the smoothing sample period which increases confidence in the housing market, thus the desire for ownership. In the upswing of the real estate cycle, households present a higher willingness to buy a housing unit either because the potential first time buyers are afraid the future house price will be so high that they may not be capable to purchase later or because households see it a good investment to gain profit in the future. This effect is magnified in the Dutch context where the highly inelastic supply leads to stronger excess demand given positive shocks. Another possible reason for such an income elasticity is the favorable tax treatment that strongly stimulates house demand, especially for higher-income groups.

The coefficient of the interest rate shows a negative effect on house prices, which is in line with previous studies by Abelson *et al.* (2005) and Agnello & Schuknecht (2011). The estimation result indicates a mortgage rates elasticity of house price of 0.31. It's worth stressing that the logarithm of mortgage interest was used in the long run equation, which means house price would change proportionally in response to the change in mortgage rates. This differs from many other studies that chose to apply the mortgage rates without taking the logarithm.[13] To compare the results with other studies, the derived mortgage rate elasticity was converted into the comparable level — a price elasticity with respect to the absolute mortgage rate — by dividing the mortgage rate, leading to a range of [−8.4, −3.0]. This consists with literature. Iossifov, Cihák, & Shanghavi (2008) summarized, for example, the international findings of such elasticity ranging from zero to around negative eight. Studies in the Dutch market have reported a variation of the impact from mortgage rates, with a range of [−0.086, −6.5]. Compared to these studies, our time-variant mortgage rate elasticity generates a median value of −4.5, though the maximum is slightly higher.

The result indicates that the extent to which households respond to the change in mortgage rates is dependent of the present magnitude of the rate itself. This provides a potential reason for the booming house prices a few years before 2008 when the mortgage rates stayed in a relatively low range. Although the debate continues whether interest rates impose a sizeable impact on house prices, our findings imply a considerable influence of mortgage rates in the Dutch housing market. Many studies have explained the channel from interest rates to house prices through the user cost framework proposed by Poterba (1984). For instance, Kuttner (2013) studied a dynamic user cost channel to connect the interest rates to house prices, apart from the credit channel that concentrates on the borrowing limit and the availability of the credit — a lowered interest rate would lead to an ease of the accessibility to credit for some households — and the risk-taking channel that indicates financial intermediaries are more willing to lend, given a lowered mortgage rate. Impacts through these mechanisms are expected to be amplified for the Dutch market, if any. The generous fiscal policy is equivalent to a further reduction in the mortgage rates, which is sizable, given the marginal tax rates in the Netherlands. This potentially leads to an initial increase of house prices to a larger extent than it should have reached in the presence of an unanticipated decrease in the mortgage rate.

Inflation exerts a significantly positive influence on house prices in our study. Given the absence of a negative association between inflation and house price, this finding runs counter to the money illusion view that households tend to rush into renting by undervaluing or ignoring the influence of inflation (Brunnermeier & Julliard, 2008). Additionally we found no evidence for the 'tilt effect of inflation' that imposes extra burden for the mortgage payments in the earlier years of the mortgage loan, which suggests a negative association between house prices and inflation (see, e.g. Modigliani & Cohn, 1979; Tucker, 1974; Kearl, 1979). The positive correlation between inflation and asset prices, however, has been reported in literature (Barber, Robertson & Scott, 1997; National & Low, 2000). Intuitively, houses, as a type of real assets, provide the possibilities to hedge against further inflation in the future (see, e.g. Barber *et al.*, 1997). On the other hand, the positive association between house price and inflation can be well explained from the user-cost view. A reasonable increase of inflation would reduce the cost of owning housing units and stimulate the housing demand. This positive impact can be further amplified when the housing market is at the upswing as households expect a higher future return of being a home-owner in general.

To further examine how inflation affects house prices, we also provide the estimation results for inflation and real house prices in Table 8. Consistent with the user-cost view, inflation still affects house prices in real term, albeit to a smaller extent.

6.3. Robustness check

We used two subsamples covering the first quarter of 1995 to the first quarter of 2008 to check the robustness of the nominal results as reported in Appendix 2. Apart from a different sample period, we also used a house-price index that partially takes the compositional change into account[14] as a different measure of house prices. The house-price index was

Table 8. Output of log(real house price) in the long run (dependent variable = $\ln P_t^r$).

Variable	Coefficient
c	−2.43***
$\ln Y_t^r$	1.70***
$\ln i_t^r$	−0.27***
π_t	0.12***

Notes: Asterisks show a different significance level, '***', '**', '*' indicate a significance level of 1%, 5%, and 10%, respectively. The output also reports an adjusted R^2 of 0.99.

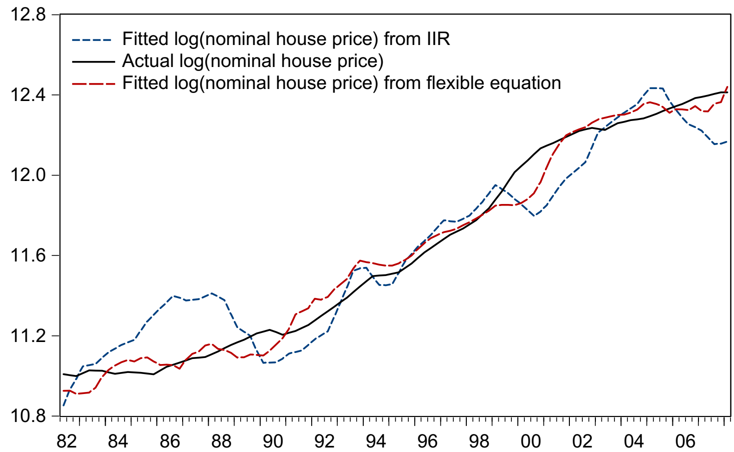

Figure 5. The gap between the log(house price) and equilibrium based on IIR and flexible equation.

obtained from the CBS and was calculated based on the housing sales for all the existing dwellings registered at the Kadaster (Dutch Land Registry Office). Both subsample estimations have the expected sign and magnitude but the inflation coefficient indicates no significant influence in the subsample estimation for nominal house prices. This might be explained by the record of a less volatile inflation during 1995–2008. These subsample estimations indicated a general robustness of the result, though not in a perfect sense. In general, we found robust results for the house price relationship in the long run.

House prices in the Netherlands have gone through various phases (see e.g. Boelhouwer *et al.*, 1996; Agnello & Schuknecht, 2011). Although we found the evidence of a long-run equilibrium among house price, household income, interest rate, and inflation, it would be interesting to investigate a more precise relationship which covers different phases in the future.

6.4. Comparison with IIR-based results

To compare with the results generated from IIR, we used the same quarterly data to re-estimate the work by de Vries & Boelhouwer (2009) who considered the IIR as the equilibrium measurement of house prices and used error correction method to estimate the house prices. Figure 5 compares the long-run house prices in logarithm generated from IIR-based estimation and that of the flexible relationship with respect to the actual prices. Our flexible relationship indicates a long-term series less volatile than that generated from the IIR-based model. In particular, the flexible model suggested that the house prices were slightly overvalued before 2008, but became undervalued in the first quarter of 2008. This difference is expected because the IIR-based model resembles precisely the interest only mortgage which usually requires small amount of capital to enter owner-occupancy, potentially leading to the long run house prices that are positively biased.

Compared to the original OTB interest-to-income ratio, the flexible long-run estimation yielde expected signs for all the variables. However, we did not find the evidence favoring the one-to-one restriction between income and interest rate as suggested by Equation (6). This does not necessarily mean an isolation between the IIR ratio and the present long-run relationship. The OTB model describes a house-price model implicitly from the perspective of affordability by proposing an interest-to-income ratio as the equilibrium. The ratio is thought to be capable to reflect the affordability of owner-occupancy in the context of the Dutch market. The current long-run equation relaxes the coefficient restrictions imbedded in the OTB model and reaches a flexible association which can be further derived from a demand-oriented model. This implies that all of the fundamentals taken into consideration are still associated with the ability to obtain housing.

We also carried out the *ex post* and *ex-ante* forecasting to compare the performance of the two models. Our strategy is to use the flexible model to forecast the house prices in the period 2005Q2–2008Q1. We did not use period that includes more recent years due to the structural changes brought in by the crisis. An advantage of using the pseudo out-of-sample forecasting is that it allows for the *ex post* forecasting, which can be further compared with the actual house prices. The two types of forecasts are presented in Figure 6 and Figure 7, respectively. The results based on the flexible model generate more accurate forecasting for both the *ex post* and *ex-ante* forecasting than that based on IIR for the nearest neighborhood that contains at least the following three periods.[15] The average *ex-ante* forecasting error generated from flexible model is 0.77% for the first subsequent three periods, which is less than 2/3 of that generated from the IIR-based model.

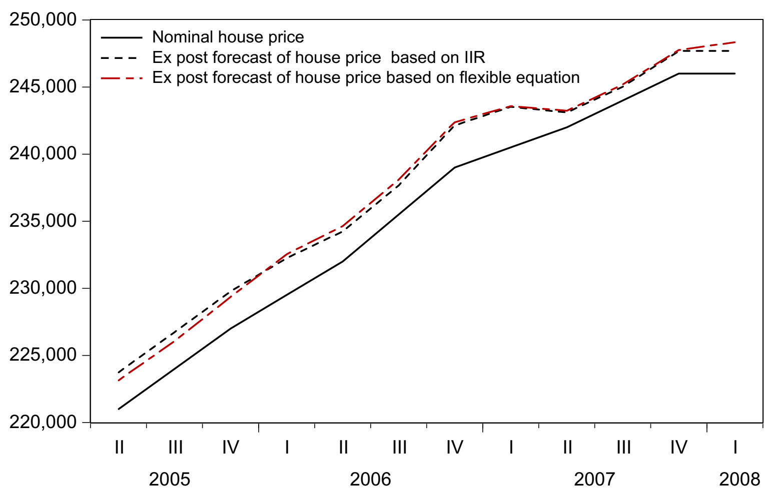

Figure 6. *Ex post* forecast of house prices (euro).

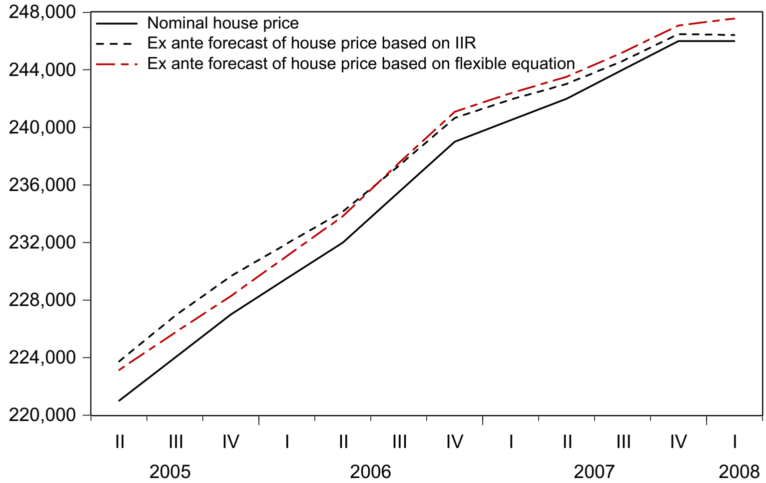

Figure 7. *Ex ante* forecast of house prices (euro).

7. Conclusion

This study investigates the long-run house prices in the highly regulated Dutch market. Special attention has been given to the existence of equilibrium and its formulation. We tested the equilibrium relation derived from a simple theoretical ground using aggregated national data and did find evidence of such an equilibrium in the Dutch market where stringent regulations are applied.

Interestingly, we found that inflation has an impact on house price in the long run, although the precise manner in which it affects the system is not yet clear. Theories interpret the influence of inflation in different ways. The asset-pricing view claims a positive impact of inflation on house prices by means of how the user-cost is determined. In contrast, the money illusion perspective suggests that increased inflation contributes to a drop in prices, as households may erroneously evaluate the value of rents by ignoring the impact of inflation, thus leaning toward a rental choice. It is possible to expect that the effect of inflation found in our study was either brought about by user-cost or as a combined consequence. This also indicates that the Dutch case, though portrayed as a regulated market, still has similarities to liberal markets. We also found that a significant impact of mortgage interest on house prices, suggesting it can be used as an effective instrument of housing market governance.

These findings seem to have some implications for policy. For example, given that inflation has functioned as an intermediary between the housing market and the macro-economy, relevant adjustments in monetary policy would transfer to housing market dynamics. Another direct and effective tool is to supervise the evolution of mortgage interest rates and the associated tax policy.

This study emphasizes a demand-oriented market and takes the perspective of household's decision-making and affordability. The setting is reasonable as the housing market is depicted as a stock market due to the small fraction of newly built houses. However, we acknowledge that the additional provision of housing might play a role over a sufficiently long term. It requires further investigation in the future. In addition, we tried to provide some clues for the recent decline housing market by investigating the periods before the crisis. Our results suggest that the house prices are overvalued to a limited extent right before

2008, indicating the decline of the market is most likely to be caused by external factors. In fact, fundamentals applied in our model showed no shocks after 2008, suggesting the decline is most likely to be caused by other external factors. It is thus expected to see a rebound of the house prices. The implication is in line with the recent report by CBS that suggests house prices in recent years seem to start a new round of boom. The period covered by the current study excluded structural changes such as crisis and policy changes, which would be interesting to investigate further.

We also acknowledge that taking a constant tax relief is a strong presumption. More insightful results are expected if data permits.

Even though this study focuses on the regulated Dutch market, it provides insights into incorporating the regulatory environment in an indirect way. In particular, regulatory elements may not be easily measured, but they can be captured to a certain extent in the model setting. This study is also connected with other regulated markets, especially in Europe, as many European countries share similar regulated features, albeit to a different extent. We expect to find similar results for those countries. Our findings may also provide some insight for other countries that move closer towards the Dutch system.

Notes

1. There is a large body of literature on the connection between the recent subprime crisis and housing market, including the work by Aalbers (2008) who discussed the channel between housing finance and mortgage crisis, and the study by Duca *et al.* (2010) who emphasized the role of financial innovation and inefficiency in the housing market to the recent crisis.
2. The range of regulated markets is broad and heterogeneous, including supply restrictions such as zoning plans and demand-associated arrangements like the tax deduction. Tu, de Haan and Boelhouwer (2016) discussed the different aspects of regulations.
3. The aim of rent control is to improve the housing condition of the low income group and to allow them to approach more decent dwellings. The process could last for years depending on the individual situation of the households. We thank an anonymous referee for this point.
4. In the Netherlands, the interest-only mortgage was developed to take advantages of the tax deduction. This type of mortgage captured a dominant share of the mortgage market; for instance, in 2006, its share of all single mortgage types was 44%.
5. We use nominal variables rather than real terms in this study partially because the impact of inflation may still exist in the long run for both nominal and real variables.
6. Stationarity here refers only to weak stationarity that requires a time-invariant property for the first two moments.
7. The estimators are superconsistent because their converging speed is faster than normal.
8. Ignoring non-stationary time series may result in a misleadingly high R^2, a relatively low Durbin–Watson statistic, and further spurious results for the reason that conventional coefficient tests become invalid for non-stationary series (Granger & Newbold, 1974).
9. A more recent test proposed by Ng and Perron (2001) (the NP test) tries to mitigate potential problems of the unit root test by using other means of estimation and modifying the lag selection.
10. We used a general inflation because the inflation excluding the housing service is not available due to data limitation.
11. This income data gives overall information of all households. We acknowledge that the income information targeted at home occupiers will be preferred if data is available. During the period of interest, the owner-occupied sector has experienced a stable increase from 42% in 1982 to more than 56% in 2008. We thank an anonymous reviewer for this point.
12. Although the max-eigenvalue test suggests the same result at 19% significance level, we chose to stick to the trace test result.
13. We found it appropriate to use the logarithm of mortgage rates, especially from the derivation of the simple model provided in the appendix.

14. Although this house price index may be superior as a measure than the average house price which neglects the heterogeneity of housing , we did not apply this index in the main estimation because the CBS only provides open records of the house price index since 1995.
15. Because forecasting error accumulates as the forecasting steps increase, the focus of comparison is given to the performance of the periods that follow subsequently the estimation period.

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Disclosure statement

No potential conflict of interest was reported by the authors.

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Appendix 1. Illustration of a demand-oriented market

We consider a representative household that determines its allocation on housing and a composite good. This household tries to get the optimal utility in the form of

$$u\left(P_t, C_t\right) = {a_1 * C_t^{1-m}}/{1-m} + {a_2 * P_t^{1-n}}/{1-n}, \tag{A1}$$

where P_t and C_t are the house price and the spending on the composite good, respectively; a_1, a_2, m, and n are preference-related parameters.
This household maximizes the multi-period utility level.

$$\sum \beta^t u\left(P_t, C_t\right), \tag{A2}$$

subject to

$$\gamma L_t + C_t = Y_t + F_t i_t L_t, \tag{A3}$$

where β is the discount factor; the mortgage loan L_t is a percentage of the house price, that is, $L_t = \theta \times P_t$; Y_t represents income; i_t is the interest rate; F_t is the tax relief factor; γ is a constant.
Equation (A3) implies that the household pays a periodic amount to repay the loan and the associated interest, and that the household also benefits from the tax deduction.
The optimal solution becomes.

$$u_p / u_c = \gamma\theta - \theta F_t i_t, \tag{A4}$$

and u_p and u_c are the partial derivatives of the utility function with respect to P_t and C_t.
We further assume the fixed periodic fraction γ is formulated in the function.

$$\gamma = f\left(F_t, i_t\right) = c_0 F_t i_t, \tag{A5}$$

where c_0 is a constant.
Combining (A4) and (A5), we can get the final solution

$$\ln P_t = c_1 + \frac{m}{n} \ln C_t - \frac{1}{n} \ln i_t, \tag{A6}$$

where constant c_1 equals $\frac{1}{n} \ln \frac{a_2}{a_1(c_0-1)\theta F}$ because the tax relief factor is treated as fixed in the Dutch context.
The second component on the right side of Equation (A6) indicates the allocation on the composite good. This amount is strongly associated with the income level. As higher income stimulates consumer demand, we assume that this household determines current spending on the composite good by considering current income and future income:

$$C_t = a_0 Y_t^{\alpha} E(Y_{t+1}^{1-\alpha}), \tag{A7}$$

where a_0 and α are parameters. Equation (A7) states that the consumption amount is determined by a weighted average of the household's income flows.
Rearrange (A7) as.

$$C_t = a_0 Y_t E\left(\frac{Y_{t+1}}{Y_t}\right)^{1-\alpha}. \tag{A8}$$

The spending on the composite good is largely determined by the expected growth in income. We separate the increase of income into two elements: the expected inflation rate (π_t^e); and the expected real income growth rate (y_t^e). Using the simple adaptive expectation form $E_t(\pi_{t+1}) = \pi_t$, this household takes the current level inflation (π_t) and income growth (y_t) as the corresponding expected values; thus, the spending on the composite good is.

$$C_t = a_0(1 + \pi_t^e + y_t^e)^{1-\alpha} Y_t, \tag{A9}$$

or equivalently,

$$\ln C_t = \ln a_0 + \ln Y_t + (1 - \alpha)\ln(1 + \pi_t + y_t). \tag{A10}$$

In the Netherlands, the increase in income shares a similar magnitude with inflation and tends to co-move with inflation. Thus, it is reasonable to assume that $y_t = a_3\pi_t$, and a_3 is constant. As $(1 + a_3)\pi_t << 1$, we approximate (A10) as.

$$\ln C_t \doteq \ln a_0 + \ln Y_t + (1 - \alpha)(1 + a_3)\pi_t. \tag{A11}$$

Combining (A6) and (A11), we reach the demand-oriented house prices in the flexible form of

$$\ln P_t = c_2 + \frac{m}{n}\ln Y_t - \frac{1}{n}\ln i_t + \frac{m(1 - \alpha)(1 + a_3)}{n}\pi_t, \tag{A12}$$

where c_2 again is the constant $c_1 + \frac{m}{n}\ln a_0$, or $\frac{1}{n}\ln\frac{a_2 a_0^m}{a_1(c_0 - 1)\theta F}$.

Appendix 2

Table A1. Subsample estimation (dependent variable = ln P_t; 1995Q1–2008Q1).

Variable	Coefficient	Std. Error	*t*-Statistic	Prob.
c	−9.58***	1.49	−6.42	0.00
ln Y_t	2.48***	0.15	16.06	0.00
ln i_t	−0.14	0.11	−1.28	0.21
π_t	0.19***	0.06	3.26	0.002

Notes: Asterisks show a different significance level, '***', '**', '*' indicate a significance level of 1, 5, and 10%, respectively. The output also reports an adjusted R^2 of 0.95 and a Durbin–Watson statistic of 0.12. The cointegration test is approved.

Table A2. Subsample estimation (dependent variable = ln(index); 1995Q1–2008Q1).

Variable	Coefficient	Std. error	*t*-Statistic	Prob.
C	−20.51***	1.60	−12.81	0.00
ln Y_t	2.75***	0.22	12.54	0.00
ln i_t	−0.21*	0.15	−1.65	0.07
π_t	0.23***	0.03	2.17	0.0006

Notes: Asterisks show a different significance level, '***', '**', '*' indicate a significance level of 1, 5, and 10%, respectively. The output also reports an adjusted R^2 of 0.95 and a Durbin–Watson statistic of 0.13. The null hypothesis of the Engle–Granger cointegration test is approved.