Permanent and transitory shocks in owner-occupied housing: A common trend model of price dynamics

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ABSTRACT

Significant fluctuations in house prices have received considerable attention in recent years. An understanding of the forces underlying the departure from fundamental values is important in explaining the mechanisms underlying housing market performance and predicting potential house price changes in the future. This study constitutes the first use of a common trend (CT) model to analyze private house prices in the Swedish market. We employ a cointegration system to analyze the macro variables of consumption expenditure per capita, user costs and house prices. We decompose shocks into those resulting from fundamental variables, specified in this research as income and the interest rate, and those resulting from cyclical variables. The results indicate that interest rates play a dominant role in explaining house price swings, and are also significant in determining user costs for households in Sweden. Transitory shocks are found to have little explanatory power for house prices and user costs in the long run. A number of tests have been performed to verify the robustness of the specification and results.

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1. Introduction

The very substantial house price fluctuations witnessed recently in many countries have created increased interest in the link between house price changes and macroeconomic conditions. House prices play a critical role in the stability of national economies and financial markets. The interaction of housing, financial and economic activities is of central importance to the performance of the economy as a whole (Muellbauer and Murphy, 2008). It is therefore important to understand the short-run dynamics involved by capturing business cycles and to understand the long-run dynamics by explaining and forecasting the macroeconomic outlook.

Two important issues in housing economics are the extent to which macroeconomic "fundamentals" are factored into the development of house prices and the extent to which these housing market activities conform to the broader business cycle. Thus, capturing the long-run forecasted economic developments related to movement in this market will contribute to a better understanding of potential house price changes in the future. Such an analysis will also inform public policy aimed at maintaining sustainable house prices and economic growth. The topics addressed herein are particularly important, and somewhat controversial, given the precarious territory the world's housing economy has found itself in since 2007.

In this study, a common trend (CT) model is constructed using Swedish housing prices. Permanent and transitory movements in user costs and house prices are estimated within a three-variable system comprising private consumption expenditure, user costs and house prices. Within this framework, the permanent components of the endogenous variables can be identified as income and interest rate shocks, based on a minimal set of identifying assumptions. These variables enable a meaningful economic interpretation of long-run house price forecasts, whereas the transitory components relate to a temporary cycle. From

the perspective of long-run movements, the present study investigates how the fundamental macro variables permeate the market to influence price and value, thus revealing the relative effectiveness of permanent and transitory shocks on market fluctuations.

The CT model has been used in macroeconomic analysis (see King et al., 1991; Mellander et al., 1992; Warne, 1991; Hjelm, 2002; Lettau and Ludvigson, 2004), but has rarely been applied to the property market.[1] The real challenge in modeling house price volatility with the CT model lies in the interpretation of structural shocks, analysis of which is based on both economic theory and the strict assumptions of the model. To the best of our knowledge, this is the first study to employ the CT model to investigate house price volatility. This method contributes to the house price literature by incorporating a more dynamic view of market equilibrium. The model treats house prices as one of a set of endogenous dependent macro variables that maintain a cointegrated relationship over time. It measures the effect of shocks that lead to transitory deviations from steady-state levels without affecting the steady state itself and that of shocks that change the steady-state equilibrium from the initial steady state to a new one, thereby yielding insights into the driving forces behind price fluctuation. It is essential that we understand the timing and magnitude of house price changes (Lettau and Ludvigson, 2004) and capture potential price movements. Although permanent shocks dominate house prices and user costs in this study, we also find a significant impact running from such costs to transitory shocks in the short run, thus shedding light on the channel through which housing prices can influence business cycle components. Hence, employing the CT model affords us comparative advantages over general cointegration studies that focus only on the trends in macro variables.

The remainder of this paper is organized as follows. In Section 2, we outline the theoretical framework of the CT model in the context of the owner-occupied housing market. We describe the data and the cointegration relationships of the variables in Section 3. In Section 4, we present our results, and in Section 5 test the stability of the model. Section 6 concludes the paper.

2. Common trend model of owner-occupied housing

In the context of a life-cycle model, households maximize their utility functions within the limits of their total household resources, interpreted as present and future income. Optimization results from an arbitrage condition, where the rate of return on housing equals the return available from alternative investments, analogous to Meen (1990) and Miles (1994), and can be expressed as:

$$\frac{U'_H(C_t, H_t)}{U'_c(C_t, H_t)} = \text{hp}_t \times \text{UC}_t = \text{hp}_t \times [(1-\theta)R_t + T_p - g_t^e], \qquad (1)$$

where U is the utility function, C_t is household consumption, H_t is housing stock, hp_t is house prices, UC_t is user

cost, which is defined as $\text{UC}_t = (1-\theta)R_t + T_p - g_t^e$, where θ = the marginal tax rate on income; R_t = the pre-tax nominal interest rate; T_p = the effective property tax rate; and g_t^e = expected price inflation. This first-order condition suggests a long-run relationship among hp_t, UC_t, C_t, and H_t.

Further for $\frac{U'_H(C_t, H_t)}{U'_C(C_t, H_t)} = f(C_t, H_t)$, if we assume that it is a log-linear function, then there would be a linear or cointegrating relationship among $\log C_t$, $\log H_t$, $\log \text{hp}_t$, and $\log(\text{UC}_t)$. Following Poterba (1984), we assume the quantity of housing units (H_t) to be a positive function of house prices in the long run, where $H_t = \varphi(\text{hp}_t)$ Poterba (1984) explains this issue on the basis of long-run price supply elasticity and the production possibility frontier between houses and other goods. If the fi-function was log-linear there would be a linear relationship between $\log C$, $\log \text{hp}$, and $\log \text{UC}$. In this study, however, we are only taking logs of C and hp, not of UC. The reason is that it is difficult to measure UC with any precision and measured values become negative for some periods.

Specifically, the endogenous variables used in the paper are:

$$x_t = [\text{CON}_t, \ \text{UC}_t, \ \text{HP}_t],$$

where CON_t ($=\log C_t$) is consumption expenditure per capita, UC_t is user cost, and HP_t ($=\log \text{hp}_t$) is house prices.

This theory allows us to employ the CT model with a minimum set of identification assumptions, and enables us to identify the permanent shocks. If there is one cointegrating relationship among three I(1) non-stationary variables, then we can identify two permanent shocks and one transitory shock, which is conditional upon the information contained within the system implied by the CT model. The cointegrating relationship of the variables requires empirical testing, which we perform in Section 4.

Compared to the general model for house prices (Painter and Redfearn, 2002; Capozza et al., 2004), we restrict our study to a small constrained system derived from the basic household life-cycle model, which enables us to further identify the shocks based on a minimal set of identification assumptions, a unique feature of this paper.

According to Eq. (4) of the CT model in the appendix, the levels of the variables are decomposed into permanent and transitory components, as follows:

$$x_t = \begin{pmatrix} \text{CON}_t \\ \text{UC}_t \\ \text{HP}_t \end{pmatrix} = x_0 + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \\ a_{31} & a_{32} \end{pmatrix} \begin{pmatrix} \Gamma_t^1 \\ \Gamma_t^2 \end{pmatrix} + \phi(L)v_t, \qquad (2)$$

where v_t is a purely transitory disturbance, and elements a_{ij} capture the long-run effect of the two permanent disturbances (Γ_t^1 and Γ_t^2) on the endogenous variables. To identify the two permanent shocks and estimate the long-run effect of each individual disturbance, we must achieve the complete identification of the six elements a_{ij}. As explained in the appendix, we need to impose a long-term consumption neutrality restriction, whereby the second permanent disturbance (Γ_t^2 is assumed to have no long-term effect on CON_t, but to be able to permanently affect UC_t and house prices HP_t, which is $a_{12} = 0$. The first permanent shock Γ_t^1, in contrast, influences the long-run behavior of consumption, user costs and house prices. This

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[1] However, critics argue that capital markets are imperfect and that credit conditions may destabilize the long-run relationship between national consumption and income (e.g., Hayashi, 1982; Jappelli and Pagano, 1989).

identifying assumption is consistent with the interpretation of the first permanent shock as an income (per capita) disturbance and the second as an interest rate disturbance. In addition, these two shocks are fundamental variables in understanding house price movement.

The first permanent shock that is identified as income is based on the long-run effects of income on consumption, user cost and house prices. In line with the permanent income hypothesis, which posits that household consumption is dependent upon the resources available to consumers over their entire lifetimes (Modigliani, 1986), we suggest a long-run relationship between disposable income and consumption. Both Campbell and Deaton (1987) and Michener (1984) demonstrate cointegration between consumption and disposable income, which is consistent with the life-cycle permanent income model.[2]

If, alternatively, we were to assume that income has no long-run effect on user costs, then we would create another order in the model, that of income and interest rate shocks. We regard this alternative as less acceptable theoretically. As its definition suggests, the user cost of housing is the marginal rate of substitution between housing services and consumption goods. It thus determines the real cost of enjoying housing services (imputed rent) under the efficient market hypothesis. An increase in income increases the marginal utility of the existing housing stock and therefore raises the UC in the long term. Nevertheless, in Section 5, we test the sensitivity of the model using two different assumptions, and find the paper's fundamental conclusions to remain robust to a change in the specification of the long-run relationship between the variables.

The second permanent shock's identification as interest rates is based on the assumption that interest rates have a long-run effect on user costs and house prices, but only a short-term effect on household consumption. The interest rate is one of the key components of user costs. It may change expected future values and returns or enhance the risk premium of housing consumption (Kearl and Mishkin, 1977), which in turn influences user costs in the long term. Interest rates' effect on consumption is difficult to predict on theoretical grounds (Deaton, 1992; Campbell and Mankiw, 1989; Cromb and Fernandez-Corugedo, 2004). It depends on the relative importance of income and substitution effects in determining how households allocate their resources over time (Cromb and Fernandez-Corugedo, 2004; Romer, 2001). Romer (2001) argues that unless the elasticity of substitution between consumption in different periods is large, interest rate increases are unlikely to bring about substantial rises in consumption and savings. Hall (1988) and Campbell and Mankiw (1989) demonstrate that consumption may be unrelated to the expected interest rate in the long run owing to household behavior. Carroll et al. (1997) discusses their weak relationship in the presence of uncertainty. Large empirical studies, such as those of Campbell and Mankiw (1989), Hall (1988), and Cromb and Fernandez-Corugedo (2004),

have proved consumption's lack of responsiveness to the interest rate. Angeloni et al. (2002) investigate European markets, and suggest that non-durable consumption is insensitive to interest rate changes in the long run. In the current study, we first estimate our CT model under the assumption that consumption is independent of interest rate shocks in the long run. As previously mentioned, we also test the model's robustness to the alternative assumption that the two have a long-run correlation.

3. Data and cointegration relationship

We estimate the CT model using quarterly data on real per capita consumption, real user costs and real house prices in Sweden from 1975 to 2009, a period that covers two complete price cycles.

The real house price index (HP) is a constant quality index constructed by Statistics Sweden (1998 = 1) deflated by the consumer price index (CPI). The values for this index are obtained by standardizing purchase prices with assessed values based on the location, size, age, and quality of the dwelling in question. We graph changes in the real price index over the study period in Fig. 1. Real house prices resemble a non-stationary process with cyclical movements: periods of price escalation are followed by bust years, as observed in the early 1980s and in 1993. These changes reflect the slowdown in real construction and sharp increase in interest rates in the 1980s and the deep economic recession and reduction in interest subsidies embedded in the tax reforms of the 1990s (see Englund et al., 1995; Hort, 1998; Barot, 2001). The most recent upward swing in real housing prices started in 1996. After 2007, house prices fell slightly, but increased again after 2009.

User cost (UC) is a derived variable measured as a percentage. It refers to the widely used real housing user cost of capital, as defined in Poterba (1984), Hort (1998) and Capozza et al. (2004) and explained in Eq. (1):

$$[(1 - \theta)R_t + T_p - g_t^e].$$

To calculate the marginal tax rate in UC, we separate the entire period into three phases: before 1982, from 1982 to 1992 and after 1992. For the period prior to 1982, income taxes are calculated by income and wealth distribution, according to the Statistics Sweden database ($\theta = \frac{\Delta t}{\Delta i}$, where t is tax liability and i is taxable income),[3] by applying the tax rules in various years to the average annual income of all employees who have declared income. From 1982 until the tax reforms of 1991, we adopt a 50% marginal tax rate.[4] Since 1992, the statutory maximum marginal tax rate has been 30% nationally, and thus a fixed rate of 30% is used for this period.

The interest rate is the five-year government bond rate. We estimate expected inflation by averaging the preceding annual changes ($\frac{\Delta CPI}{CPI}$). We obtain the government bond rates and price level data from Riksbank of Sweden.

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[2] However, critics argue that capital markets are imperfect and that credit conditions may destabilize the long-run relationship between national consumption and income (e.g., Hayashi, 1982; Jappelli and Pagano, 1989).

[3] See Englund et al. (1995) for a more detailed discussion.
[4] See Kanis and Barot (1993) for a more detailed discussion.

$$g_t^e = (P_t - P_{t-4})/P_t.$$

The consumption series used in this study is total price-constant consumption expenditure per capita. Total private consumption includes durable goods and is expressed in 1998 prices. Consumption per capita is total consumption divided by total population. As is generally known, the permanent income assumption incorporates the consumption of non-durable goods and the service flow from the consumption of durable goods. Unfortunately, such service flow for housing is implicit and difficult to value. As an alternative, we could employ total consumption excluding durable consumption, but doing so could lead to other problems, such as the large wealth elasticity of demand. Accordingly, it is not adopted here.[5] All three types of data used in this study are plotted in Fig. 1. We take the logarithm for HP and CON, but not UC in the test, as previously noted.

To test the long-term relationships among the three variables, we employ the standard Johansen maximum likelihood method. Adopting the Akaike Information Criterion (AIC) and log criterion (LC), we find use of a linear deterministic trend with a four-period lag to produce the best fit. The cointegration results are reported in Table 1, along with the unit tests for each variable. These tests confirm the presence of one cointegrating vector for the variables, which suggests two common trends in the system. In Section 5, we test the stability of cointegration using the method proposed by Hansen and Johansen (1999).

4. Empirical results

In the previous section, we reported one cointegration relationship for the vector $x_t = [CON_t, UC_t, HP_t]$. Hence, the system is driven by two common trends and one temporary shock. Theoretically, we can identify these two permanent shocks as an income shock and an interest rate shock. In this section, we present variance decomposition analysis and the impulse response functions, which can in turn test our assumptions with respect to the empirical results. A dummy variable is also included in our CT model to adjust for seasonal effects.

The following is the estimated CT model. All of the effects of the permanent shocks are significant with the expected signs at 90%. The asymptotic standard deviations of the coefficients are in parentheses.

$$
\begin{bmatrix} CON_t \\ UC_t \\ HP_t \end{bmatrix} = X_0 + \begin{bmatrix} 0.0088 & 0 \\ (0.0023) & (-) \\ 0.01253 & 0.0251 \\ (0.0027) & (0.0099) \\ 0.0403 & -0.0411 \\ (0.00268) & (0.00162) \end{bmatrix} \begin{bmatrix} T_{inc,t} \\ T_{in,t} \end{bmatrix} + C(L)^* \varepsilon_t
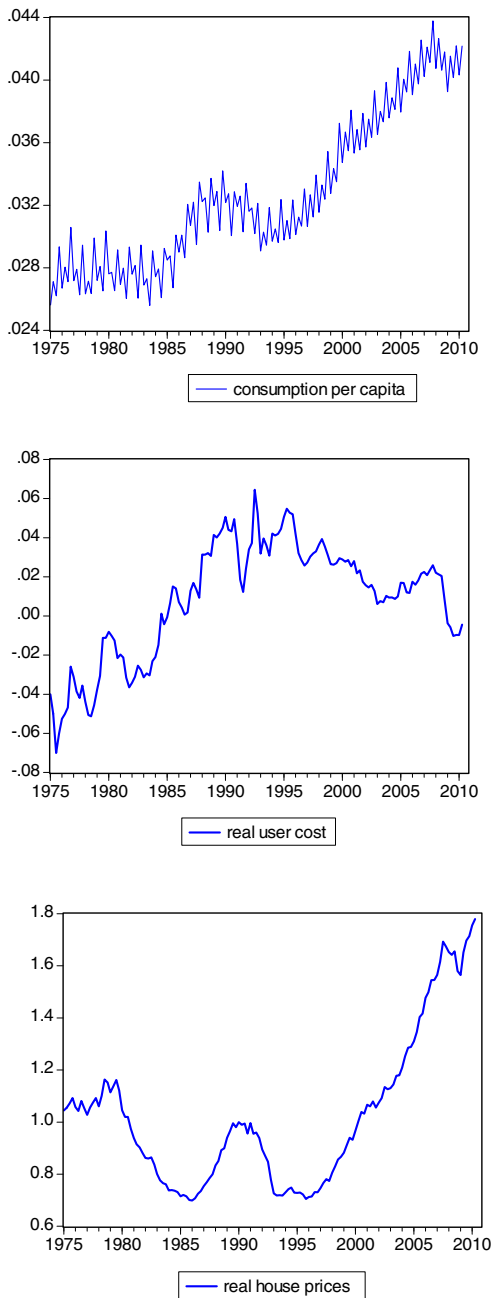$$

(3)

Fig. 1. Three Time series used in the study. Quarterly data are used for all series from 1971 to 2009. Consumption is ratio between total price – constant consumption expenditure divided by total population. User cost is reduced variables from the paper and is measured in percents. House price index (HP) is a constant quality index. More on data information can be found in Section 3 in the paper.

Several empirical proxies are used to measure house price appreciation, including that based on the CPI (see Capozza et al., 2004; Hort, 1998) or actual house price inflation and forward- or backward-looking expectation (see Miles, 1994; Meen, 2002). Here, following Miles (1994), we employ static backward-looking expectation, which is measured as the rate of annual changes in nominal housing prices:

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[5] A number of authors, such as Case et al. (2005), Rudd and Whelan (2002) and Koop et al. (2005), argue that it is total consumption expenditure that matters with regard to the intertemporal budget constraints on spending. They also question the validity of using the consumption of nondurable goods in studying housing wealth.

Table 1
Cointegration tests and unit roots tests.

Cointegration test				
Included observations: 137 after adjustments Trend assumption: Linear deterministic trend Series: LOG(CON) UC LOG(HP) Lags interval (in first differences): 1–4				
Hypothesized No. of CE(s)	Eigenvalue	Trace statistic	0.05 Critical value	Prob.[**]
Unrestricted cointegration rank test (Trace)				
None[*]	0.153166	32.05415	29.79707	0.0270
At most 1	0.055404	9.277799	15.49471	0.3403
At most 2	0.010666	1.469077	3.841466	0.2255
Variables	Level no trend		First difference no trend	
Unit roots tests				
Consumption per Capita (CON)				
ADF test (8 lag)	−0.024 (−2.88)		−3.35	
PP test	−2.46		−48.49	
User cost (UC)				
ADF test (0 lag)	−1.86		−10.32	
PP test	−1.86		−10.31	
House price (HP)				
ADF test (8 lags)	−1.52		−3.12	
PP test	−0.006		−3.12	

Trace test indicates 1 cointegrating Eqs. at the 0.05 level.
Note: critical value to reject null hypothesis at 5% significance level is −2.88.
Lag value for ADF test are based on SCI.
Natural logs of the CON and HP are taken.
  [*] Denotes rejection of the hypothesis at the 0.05 level.
  [**] *p*-values.

4.1. Variance decomposition

In Table 2, we present the forecast error variance decompositions for house prices and user costs to illustrate the percentage of such variance that is attributable to innovations in the common stochastic trends for each year. With regard to the variations in house prices, the results indicate that interest rates and income are the dominant explanations for price changes, and tend to produce a stable fraction of price variation in both the short and long run. House prices can be explained primarily by the fundamental factors of income and interest rates, whereas transitory shocks have little explanatory power.

Permanent shocks from interest rates account for 72% to 94% of the long-run variations in house prices, whereas income shocks explain just 5–6%. The contributions of both variables are stable over the long run. Transitory shocks have significant explanatory power for single-period price changes in a quarter, but their contribution declines rapidly and has no explanatory power in the long run. It is important to remember that in our model, changes in credit availability are captured by housing prices and interest rates.

With regard to fluctuations in user costs, the impact of interest rates is similar to that observed for house prices. Interest rate innovations explain about 90% of user cost changes in the long run. As was the case for house prices, the influence of transitory shocks is significant in a single-period window and then terminates. However, in the case of user costs, this influence declines more slowly, with

Table 2
Forecast error variance decomposition.

	Income	Interest rate	Transitory
Real house prices			
1 quarter	0.068	0.826	0.106
1 year	0.045	0.917	0.038
2 year	0.054	0.937	0.009
3 year	0.055	0.940	0.005
4 year	0.055	0.942	0.003
5 year	0.055	0.943	0.003
6 year	0.054	0.943	0.002
Real User Cost			
1 quarter	0.056	0.819	0.125
1 year	0.056	0.789	0.155
2 year	0.049	0.812	0.139
3 year	0.046	0.819	0.135
4 year	0.045	0.902	0.053
5 year	0.045	0.914	0.041
6 year	0.044	0.922	0.033

Results of decomposition are shown for six-year time horizon for both real house prices and real user cost. It determines what proportions of changes (in percentage points) in the real house prices and user cost can be attributed to performance of two permanent shocks (income and interest rate) and one transitory shocks, respectively.

bubbles explaining about 6.5% of user cost innovations in the long run.

Finally, we find that more than 98% of the variation in consumption (results not shown here) can be explained by income in the long run, with its contribution remaining rather stable over the entire study period.

4.2. Permanent shocks: fundamental effects

The main focus of this study is analysis of the impact of permanent and transitory shocks in the housing market, with dummy variables for seasonal adjustments. However, a graphic presentation of the results for consumption can help us to interpret household behavior and confirm the robustness of our model. Fig. 2 displays the impulse response functions for consumption, user costs and house prices over an eight-year period, separately measuring the responses to income and interest rate shocks. We in-

clude consumption in this analysis to test the robustness of our model.

Income's impact on consumption is insignificant in the first 2 years shown in Fig. 2, but becomes significant thereafter. This finding may result from our use of total consumption, which includes durable goods. Because purchases of durable goods are relatively infrequent, their inclusion in the metric may render it less sensitive to income shocks. However, given the importance of household wealth as a determinant of total consumption in Sweden, and given that most households are bound by credit
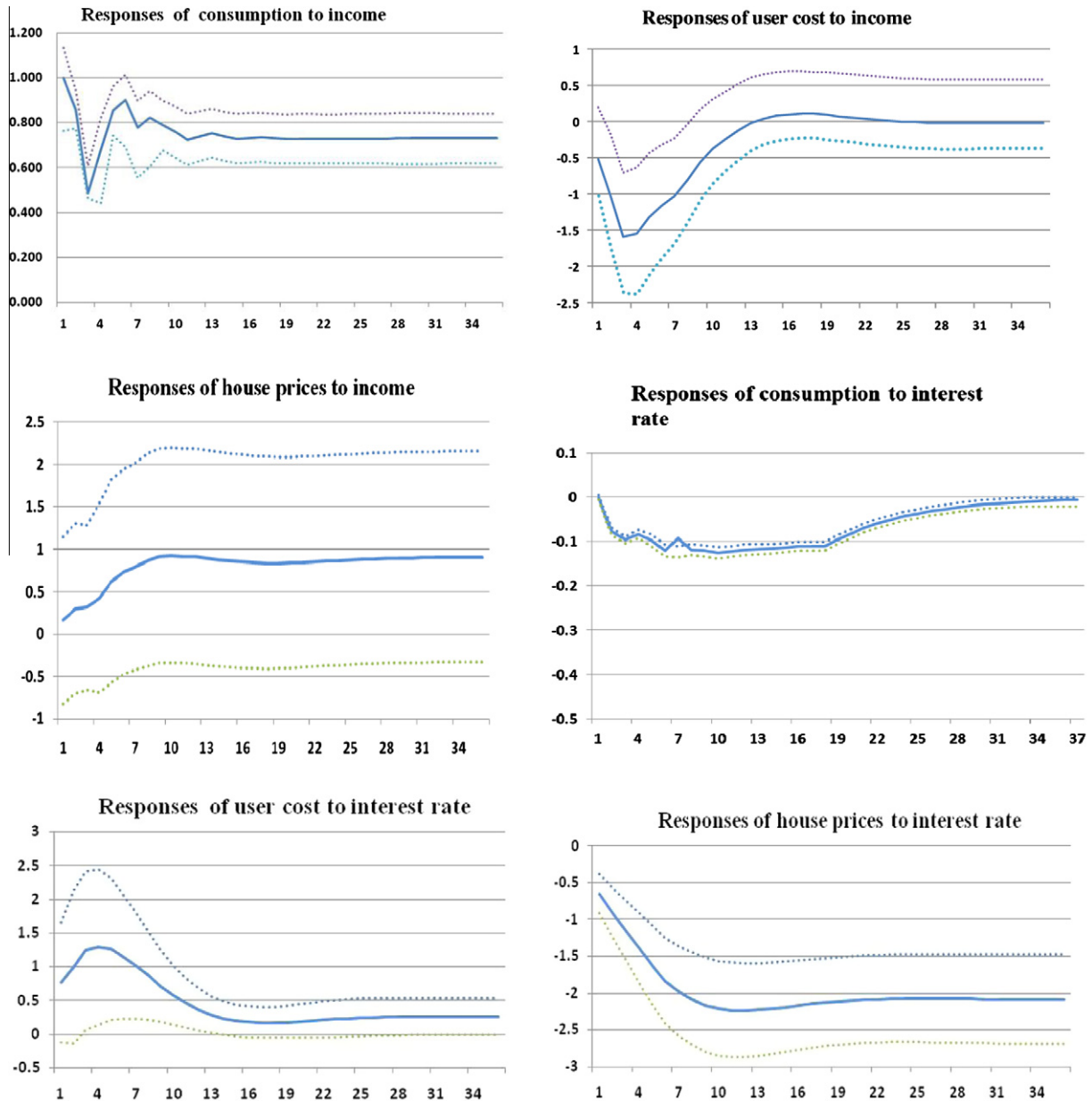

Fig. 2. Impulse response function of consumption per capita, user cost and house price on permanent shocks. The figures give the overall estimated impulse response functions associated with unitary permanent shocks in income and interest rate with one standard deviation shown in Eq. (2) respectively for 8 years. The horizontal axis denotes quarters ahead and the vertical axis denotes percent changes since HP and CON are taken log and UC is measured in percentage. The dashed lines indicate confidence bands.
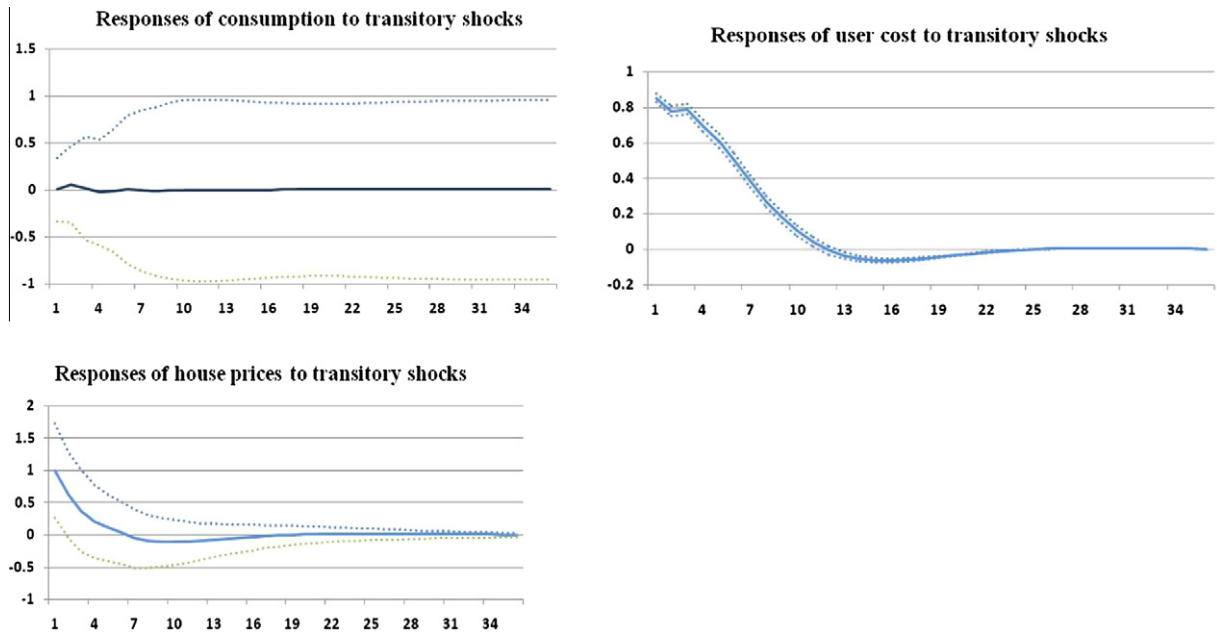
Fig. 3. Impulse response function of consumption per capita, user cost and house prices on a transitory shock. The figure gives the overall estimated impulse response functions associated with unit of transitory shocks for 8 years. The horizontal axis denotes quarters ahead and the vertical axis denotes percent changes. The dashed lines indicate confidence bands.

constraints,[6] it may be true that consumption in this country is relatively insensitive to income shocks in the short run.

User costs' response to permanent shocks in income is significant and positive in the long run, but not in the short run. Income shocks are significantly and positively related to house prices over the entire period, whereas interest rate shocks are significantly related to house prices in the short run, but with persistent negative effects. When house prices and user costs respond to changes in economic fundamentals, it takes about 3 years for them to establish the revised equilibrium levels fully.

A one standard deviation increase in interest rates is associated with a greater than 2% decline in real house prices in the long run. Prices reach their revised equilibrium level in approximately six quarters. In the long run, house price movement may be reduced to a certain extent by altering expectations of returns from new construction, such that potential price decreases may be attenuated as a result of elasticity on the supply side.

Interest rate increases imply more expensive owner-occupied housing services and a higher cost of living. We thus expect such increases to be positively related to user costs, a theoretical expectation that the results presented in Fig. 2 suggest is the reality. Interest rate shocks have a strong and enduring effect on user costs, which take about 4 years to establish their revised equilibrium level in full.

It is noteworthy that the house price response shown in Fig. 2 is likely to overshoot the equilibrium response to demand shocks in the short run, as house prices decline sharply at first and then rise gradually to the revised equilibrium, possibly because the housing stock is fixed in the short run. Such overshooting should not be identified as a bubble, but rather only as evidence of market inefficiency.

4.3. Transitory shocks: cyclical effect

Turning to transitory shocks, in Fig. 3 we plot the impulse response functions for consumption, user costs and house prices. User costs' contemporaneous responses to the transitory shocks are about 0.8%, and, accordingly, house prices increase by only roughly 0.01%. The effect of transitory shocks on user costs declines to zero with a relatively long horizon of more than 6 years.

It is difficult to interpret transitory factors, as they can comprise numerous temporary cyclical effects. They may stem from the endogenous cycles, as the endogenous variables of CON, UC and HP all produce short-run deviations from the long-run upward trend. For example, the short-term cyclical behavior of house prices is generally believed to reflect delayed supply responses to changes in effective demand. A lag response of supply could initially cause price overshoot and further lead to surplus supply. Prices tend to remain at their new equilibrium until the increased stock again catches up with demand. It is also possible that transitory shocks originate in the exogenous cycles, such as those generated by changes in monetary policy or construction supply. In addition, speculative bubbles and

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[6] According to Sweden's Riksbank (2006), "the loans have been obtained largely to finance house purchase[s], [of] which more than 85% are secured with real estate. Consequently, the amount of debt has grown relative to disposable income significantly, the debt ratio to disposable income in year 2002 is around 100%, and it arrived at 145% by the end of year 2006," which constitutes evidences in support of this supposition.
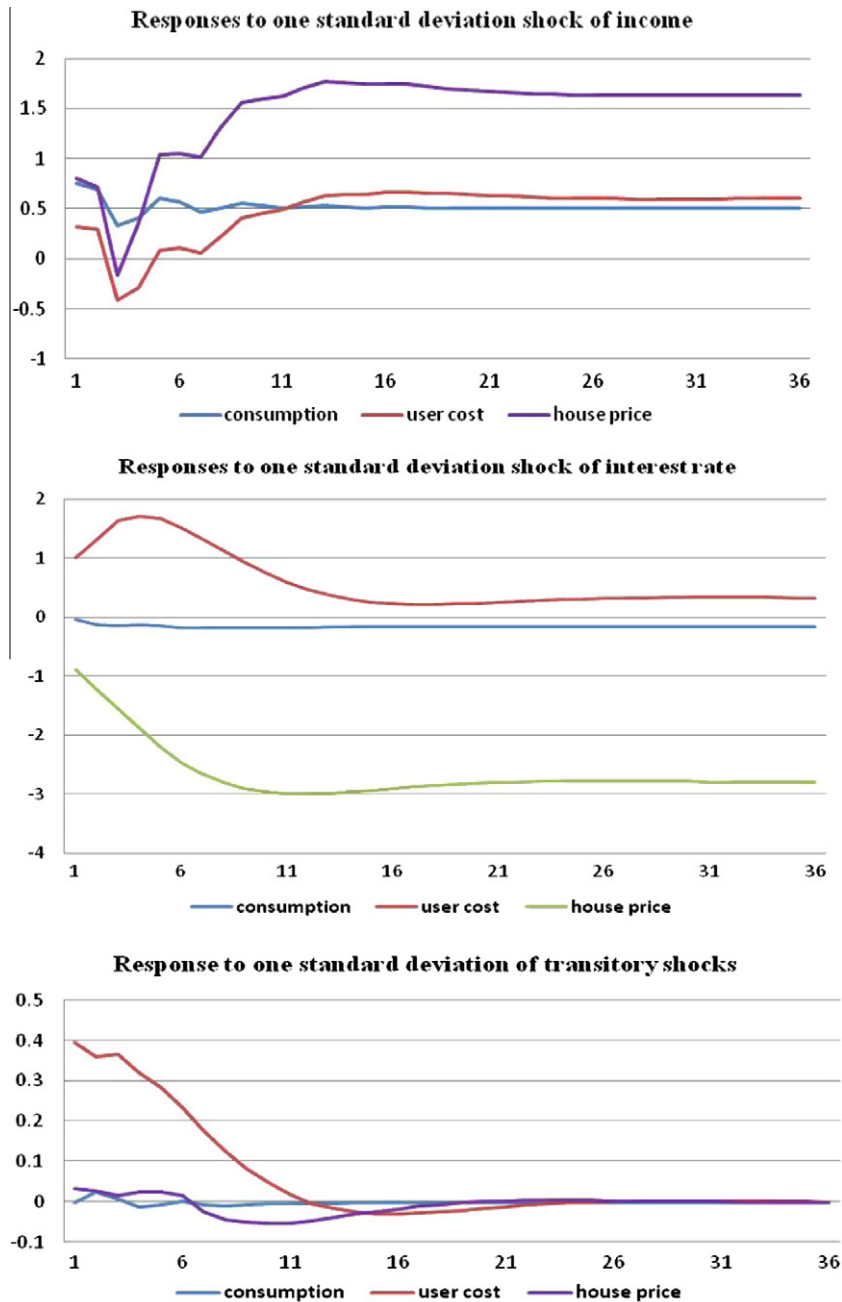
Responses to one standard deviation shock of income

Responses to one standard deviation shock of interest rate

Response to one standard deviation of transitory shocks

Fig. 4. Impulse response function of real consumption per capita, real user cost and real house price on permanent shocks and transitory shocks. (with the assumption that income has no long run effect on user cost).

seasonal fluctuations can both have temporary effects on the system.

Although we do not identify the transitory shocks in this study, we can further examine their characteristics simply by examining their link with the price cycle. We estimate the transitory shocks for both the price bust (1979–1985 and 1991–1995) and non-price bust periods (1971–1978, 1986–1990 and 1996–2009). We establish the cumulative shock value during the price bust periods to be −0.025, whereas that during the non-bust periods

is 0.40. The figures in these periods are significantly different,[7] which implies a link between property cycles and transitory shocks, with these shocks performing asym-

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[7] The significance of the value difference can be tested with t-tests, as in Hjelm (2002). Such a test could be $t_{v_1+v_2} = \frac{(\bar{X}_1 - \bar{X}_2)}{S_p \sqrt{\frac{1}{n_1} + \frac{1}{n_2}}}$, where $S_p = \sqrt{\frac{(n_1-1)S_1^2 + (n_2-1)S_2^2}{n_1 + n_2 - 2}}$, $\bar{X}_i$ is the mean value, and n_i is the number of observations, $i = 1, 2$. The critical value for the 5% significance level is 1.96. Therefore, the difference in the mean value (0.42) is significant.

metrically over the property cycle peaks and troughs. Other techniques, such as the nonlinear stochastic trend model (e.g., Hamilton, 1989), could be employed to further examine whether these short-run variations depend on the state of the property cycle.

5. Stability tests of the model

The reliability of the results presented in the previous section is critically dependent on two essential assumptions of the CT model: (1) that there is one, and only one, cointegration relationship among the endogenous variables, and (2) that interest rates and consumption are fundamentally independent of each another in the long run. In this section, we test the sensitivity of our empirical approach to these assumptions.

We first test the stability of the cointegration relationship among consumption, user costs and house prices. We do so by conducting several tests of cointegration models, as introduced by Hansen and Johansen (1999). We present fluctuation tests of the eigenvalues, and also test the parameter constancy of our cointegration models.

Typically, parameter constancy is tested by analyzing the sample path of the elements in the estimated impact matrix $\Pi = \alpha \beta'$, where α is the vector of the adjustment parameters and β is a matrix of the estimated cointegrating vectors. However, Hansen and Johansen (1999) argue that a fluctuation test of Π could be biased because of the test's distribution (see Quintos, 1995) and the asymptotic distributions of α and β (see Quintos, 1997). Therefore, instead of testing the elements of Π, Hansen and Johansen (1999) test the estimated eigenvalues, λ_i, $(i = 1, \ldots, r)$, in which changes in α and β are reflected. Based on the general trend in the changes in these eigenvalues, a further LM-type test for the constancy of β is conducted using the Nyblom statistic.

For our cointegration space, the test results for the fluctuation in the eigenvalues of the endogenous variables are as follows.

Consumption	$\tau_T^{(t)} = 16.6371$	*p* value = 0.0101
User costs	$\tau_T^{(t)} = 20.4720$	*p* value = 0.0101
House prices	$\tau_T^{(t)} = 6.1215$	*p* value = 0.1414
Joint test	$\tau_T^{(t)} = 7.1402$	*p* value = 0.0505

These results suggest that parameter constancy cannot be rejected for any of the variables other than house prices. In addition, the joint test of the constant eigenvalues is accepted at the 5% level of confidence.

Because the estimated eigenvalues do not yield sufficient information on parameter constancy (Hansen and Johansen, 1999), we further test such constancy with an LM-type test, which produces the following results.

5% Critical value	
$\sup Q_t^T = 4.738$	3.69
$\text{mean} Q_T^t = 1.85$	1.32

The results of both tests are consistent with the hypothesis of constancy for the estimated cointegration parameters.

We next test the sensitivity of our findings to the assumption of long-run independence between interest rates and consumption. We do so by relaxing this assumption in the CT model's loading matrix, substituting the alternative assumption of independence between income and user costs. Having reordered the endogenous variables in this manner, we then recalculate the variance decomposition and impulse responses for the new CT, reporting the results in Fig. 4. These results are highly consistent with those generated by the unmodified CT model. The income innovations in the revised model are more dynamically related to house prices and user costs, but the fundamental findings remain the same. It can thus be concluded that the findings presented herein are robust, and not critically sensitive to the specification chosen for the CT model.

6. Concluding remarks

In this study, we apply a CT model to the private housing market in Sweden to identify the permanent and transitory shocks in that market. Guided by the life-cycle model, we construct a cointegration system comprising the fundamental factors of per capita consumption expenditure, user costs and house prices. Our test for the order of integration suggests that the model is cointegrated at order 1. Accordingly, the system can be decomposed into permanent income and interest rate shocks and transitory shocks. This distinction is based on economic theory and is expressed in the specification of the model. The results of this study indicate that interest rates play a dominant role in explaining housing price swings and are also significant in determining household user costs.

Although our explanatory results for Swedish house prices as similar to those of several previous studies, we apply the CT model to further capture the effects and channels through which different shocks affect such prices. Doing so is important for understanding natural house prices, and lays the foundation for forecasting trends in house prices.

The CT model applied in this study is linear. It thus implicitly assumes that changes in the causal variables result in symmetrical economic fluctuations. In other words, positive and negative deviations are restricted to having the same effects. Considering the interesting results obtained by Hamilton (1989), who finds the business cycle to be characterized by either positive (expansion) or negative (recession) growth, it would be useful to develop a nonlinear stochastic trend model using a Markov switching process to capture the potentially asymmetrical impacts of macro variables in different phases of the Swedish business cycle.

Appendix A

A.1. Econometric framework

The CT model is used to decompose permanent and transitory shocks in the context of a cointegration model

that considers short-term fluctuations from the perspective of long-run equilibrium.

Following Warne (1991), the structural CT model can be described as follows:

$$X_t = X_0 + \Upsilon\tau_t + \Phi(L)v_t, \tag{4}$$

where X_t is an n-dimensional vector of dependent variables, and $v_t(n \times 1)$ is white noise with $E(v_t) = 0$ and $E(v_t v_t') = I_n$. The polynomial $\Phi(L) = \sum_{i=0}^{\infty} \Phi_i L^i$ is finite. Hence, $\Phi(L)v_t$ is stationary. X_0 is also stationary, and τ_t is a vector of trends modeled as random walks with a drift:

$$\tau_t = \mu + \tau_{t-1} + \varphi_t. \tag{5}$$

$\varphi_t(k \times 1)$ is the disturbance sequence, with $E(\varphi_t) = 0$ and $E(\varphi_t \varphi_t') = I_k$. It measures the shocks that permanently affect future values of X_t. k is the number of common stochastic trends, and $k \leqslant n$. γ is the $n \times k$ loading matrix for τ_t.

We thus have

$$X_t = X_0 + \Upsilon\left[\tau_0 + \mu t + \sum_{j=1}^{t}\varphi_j\right] + \Phi(L)v_t. \tag{6}$$

Eq. (6) can now be decomposed into a permanent component and a stationary temporary component, as follows:

$$X_t^p = \Upsilon\left[\tau_0 + \mu t + \sum_{j=1}^{t}\varphi_j\right] \quad \text{is the permanent component,}$$

and

$$X_t^s = X_0 + \Phi(L)v_t \quad \text{is the transitory component.}$$

Note that v_t and φ_t are correlated in that trend innovations may exert both permanent and temporary effects on X_t. Thus, v_t is white noise caused by the permanent effects of trend shocks φ_t and transitory shocks σ_t, independent of φ_t.

To specify values for the CT model, we need to construct a reduced form. Now, assuming that $\{x_t\}$ is cointegrated of order (1,1) implies that there is an $n \times r$ matrix t a' that is orthogonal to the loading matrix Y. $r = n - k$, which equals the number of linearly independent vectors that are orthogonal to the loading matrix. Thus, we can define $Z_t := a'x_t$, which is jointly stationary.

Using the Granger Representation Theorem (see Engle and Granger, 1987), we can formulate a vector autoregressive (VAR) model,

$$A(L)x_t = \rho + \varepsilon_t, \tag{7}$$

using the following world vector moving average representation.

$$\Delta x_t = \delta + C(L)\varepsilon_t, \tag{8}$$

where $C(L) = I_n + \sum_{i=1}^{\infty} C_j L^j$ and $\sum_{i=1}^{\infty}|C_j| < \infty$.

By means of matrix algebra, we can find another polynomial $C*(L)$, such that $C(L) = C(1) + (1 - L)C*(L)$ and $C(1) = \sum_{i=0}^{\infty} C_i$, $C*(L) = -\sum_{j=i+1}^{\infty} C_j$ for $i \geqslant 0$. $C*(L)$ is absolutely summable. $\{\varepsilon_t\}$ is white noise, with $E(\varepsilon_t) = 0$ and $E(\varepsilon_t \varepsilon_t') = \Sigma$, a positive definite matrix.

If we substitute the expression for C(L) recursively into Eq. (8), then we obtain the following reduced-form CT model.

$$X_t = C(1)\xi_t + C(L)^*\varepsilon_t = X_0 + C(1)\left(\xi_0 + \rho t + \sum_{j=1}^{t}\varepsilon_j\right) + C(L)^*\varepsilon_t, \tag{9}$$

where $\xi_t = \rho + \xi_{t-1} + \varepsilon_t$ and $\delta = C(1)\rho$.

The reduced-form CT is conditioned upon the stationarity of $\{x_t\}$, which implies that $\alpha'C(1) = 0$.

Permanent shocks exert a long-run impact on the future values of the variables, whereas the impulse responses with respect to the transitory shocks decline to zero or *gradually disappear* with an increasing time horizon. According to Gonzalo and Ng (2001), the shock φ_t is said to be permanent if $Lim_{h\to\infty}\partial E_t(X_{t+h})/\partial\varphi_t \neq 0$. Analogously, shocks σ_t are said to be transitory if $Lim_{h\to\infty}\partial E_t(X_{t+h})/\partial\sigma_t = 0$.

Comparing the structures of the CT Eq. (6) and reduced CT models Eq. (9), we can identify them with $\Upsilon\varphi_t = C(1)\varepsilon_t$ and $\Upsilon\Upsilon' = C(1)\Sigma C(1)'$, $\Upsilon\mu = C(1)\rho$. Hence, we can estimate the CT model and shocks (permanent and transitory) as conditioned upon the values of C(1) and Σ, which can be derived by inverting a vector autoregressive model, as proposed by Stock and Watson (1988), or inverting the restricted VAR model, as proposed by Campbell and Shiller (1988) and Warne (1991). In this study, we follow the latter alternative and estimate the restricted VAR model:

$$B(L)y_t = \theta + \eta_t, \tag{10}$$

where $B(L) := M[A^*(\lambda)M^{-1}D(\lambda) + \gamma^*\lambda]$ and $y_t = D_\perp(L)Mx_t$.

$M = \begin{bmatrix} S_k' \\ \alpha \end{bmatrix}$ is a nonsingular matrix; S_k is a $k \times n$ selection matrix, where $S_{i,k}C(1) \neq 0$ for all $i \in \{1, \ldots, k\}$. $\gamma^* = [0 \quad \gamma)$; and $D(\lambda)$ is an $n \times n$ matrix, such that:

$$D(\lambda) = \begin{bmatrix} I_k & 0 \\ 0 & (1-\lambda)I_r \end{bmatrix} \text{ and } D_\perp(\lambda) = \begin{bmatrix} (1-\lambda)I_k & 0 \\ 0 & I_r \end{bmatrix}.$$

Hence, $C(1) = M^{-1}D(1)B(1)^{-1}M$. $\tag{11}$

It should be noted that there are two essential criteria for characterizing shocks: the cointegration relationships among the variables of interest and the orthogonal relationships among the shocks.

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