A Retrospective Analysis of the House Price Macro-Relationship in the United States*

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This study provides empirical evidence on the strengthening of the impact of house prices on the U.S. macroeconomy. The stability of the house price macro-link is tested in a small-dimensional vector autoregressive model over the last fifty years. The estimated break points are used to split the sample into different segments, and a multivariate time-series analysis is performed within sub-samples. The paper finds a robust structural break in the mid-80s. In addition, time-series analysis across segments provides evidence that the effect of house prices, not only on private consumption but also on economic activity, has intensified since the mid-80s.

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1. Introduction

A decade ago, Alan Greenspan (2002) suggested in his testimony to the Joint Economic Committee of the U.S. Congress that the strong appreciation of house prices after the crash of the stock market in 2001 might have helped save the U.S. economy from a more serious

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recession, a point of view shared by most of the economists at that time. A few years later, a large crash in housing prices triggered—in all likelihood—a financial and economic crisis. Inevitably, both events have contributed to the consensus among economic practitioners and many researchers that fluctuations in housing prices are a key factor in the U.S. business cycle.

The question of the extent to which the deep transformation of mortgage markets in recent years has modified the house price macro-relationship arose naturally. A consensus rapidly emerged in the literature that the effect of house prices on the macroeconomy has intensified since the mid-80s, a time when important deregulations took place in the mortgage market along with the development of secondary markets and the increasing role of the government-sponsored enterprises (GSEs) in mortgage financing.[1]

Consequentially, a persistent fall in mortgage transaction costs occurred along with a strong relaxation of credit conditions. Homeowners were allowed to easily extract equity from their houses and loan-to-value ratios associated with mortgage loans rose, increasing the ability of landlords to borrow against their real estate.[2]

In the literature, theoretical models showing how a relaxation of credit constraints on the mortgage sector amplifies the effect of housing prices on the macroeconomy have been developed by Miles (1992), Kiyotaki and Moore (1997), Iacoviello (2005), and Iacoviello and Neri (2010), among others. The primary mechanism of these models is that a reduction of credit constraints allows economic agents to increase the amount that they are able to borrow against their home equity and use these additional resources to finance extra consumption and investment. Thus, because of the multiple signs of a relaxation in credit conditions in the United States, we can expect that the impact of house prices on the macroeconomy has become stronger over time.

[1] See Gerardi, Rosen, and Willen (2010) for a brief review on the modifications of the U.S. mortgage market in the last forty years.

[2] Taking data from the Federal Home Loan Mortgage Corporation, we calculate a substantial fall in fees and points for a regular thirty-year fixed-rate mortgage. We obtain an average of 2.3 for the 1980s, 1.5 for the 1990s, and 0.7 for the 2000s. Greenspan and Kennedy (2008) find a sustained increase in both their gross and net equity-extraction measures (relative to disposable income) since the beginning of the 1990s.

The important role of housing prices in the U.S. business cycle in recent years does not provide sufficient evidence to conclude that the house price effect in the macroeconomy has strengthened. Indeed, one should take into consideration the magnitude of the last boom and bust in house prices; according to the Case-Shiller Home Price Index, the boom and bust in house prices has been the largest at least since the end of the nineteenth century. Although time-series and micro-econometric evidence largely show that housing prices have a positive impact on the U.S. economy (see, for example, Green 1997; Case, Shiller, and Quigley 2005; Haurin and Rosenthal 2006; Bostic, Gabriel, and Painter 2009; and Miller, Peng, and Sklarz 2011), only a few papers provide empirical evidence on the strengthening of the effect of house prices on the macroeconomy over time. Muellbauer (2008), using a structural single-equation estimation, finds that before the U.S. credit market liberalization in the 1980s, there was no significant housing wealth effect on consumption; however, since then, the effect has become positive. Altissimo et al. (2005) estimate a marginal propensity to consume (MPC) out of wealth from aggregate U.S. data. Their results suggest that the MPC appears not to have statistically changed since the beginning of the 1970s. They find, however, a big increase in the MPC at the end of the 1960s.

In this paper, we provide empirical evidence of the strengthening of impact of house prices on the U.S. macro-variables. We proceed in two steps. First, we test the stability of the house price macro-link in the United States during the last fifty years. A recursive vector autoregressive (VAR) model including gross domestic product (GDP) or private consumption, interest rates, and house prices is chosen to account for the house price macro-relationship. The methodology developed in Qu and Perron (2007) is applied to identify breaks in our multivariate system. In this method, the dates and the number of breaks are entirely determined by the algorithm, rather than being imposed ex ante like most of the methods used in the time-series literature.[3] Thus, this methodology does not require us to set a priori the mid-80s—a period systematically found as a

[3]Bai and Perron (1998) also develop a methodology to detect multiple breaks occurring at unknown dates, but the method only considers the case of a single linear equation.

structural break in the literature on the "Great Moderation"—as a unique break point in the house price macro-link. Indeed, a rupture in this relationship could have occurred as early as the end of the 1960s, when the privatization of the Federal National Mortgage Association ("Fannie Mae") in 1968 and the creation of the Federal Home Loan Mortgage Corporation ("Freddie Mac") in 1970 established the foundation of what would become the renovated U.S. mortgage market. A break might also be found as recently as the early 2000s, when a deepening of markets for securitized contracts and derivatives occurred and sub-prime mortgage loans developed at a very strong pace. Furthermore, the method allows us to consider the fact that the reduction in credit constraints has been a gradual process and that the modification in the house price macro-link did not occur all at once. The implication of this hypothesis is that the aforeproposed break dates are not exclusive and that more than one break could exist in the link. In the second step, we use the break points estimated to split the sample into different segments, and on each sub-sample, multivariate time-series analysis, including impulse response and Granger causality tests, are performed.

Our results can be summarized as follows: the application of Qu and Perron's methodology in a VAR with linear restrictions concludes that there are two break points in the house price macro-relationship when GDP is introduced in the VAR specification: one at the end of the 1960s and the second in the mid-80s, the latest rupture being the strongest among the two break dates. We also find that there is only one structural break in the link when private consumption is included in the multivariate system, in the mid-80s. Time-series analysis across sub-samples finds that the overall impact of house prices, not only on private consumption but also on economic activity, has intensified since the mid-80s. Our results give support to the hypothesis that the modification of the real-estate market in the mid-80s seems to have been a key factor in the strengthening of the effect of house prices on the macroeconomy.[4]

[4]Generally, the "Great Moderation" literature finds reduced responses of GDP to different shocks in the period after the mid-80s (see Stock and Watson 2003 for a survey), not increased responses as in this paper. However, the papers related to that literature do not provide evidence of a change on the effect of house price shocks on the economy.

The paper is organized as follows. In section 2, we present Qu and Perron's methodology, which is used to detect breaks. Section 3 applies this methodology to a small-dimensional VAR model using U.S. macro-data. In section 4, we split the initial sample into different segments using the identified break points. We then perform impulse response and Granger causality tests in each of the segments from the initial VAR model specification. Section 5 contains the concluding remarks and proposes avenues of future research.

2. Testing Structural Changes in a Multivariate Model

The various methods of identifying structural changes in a single regression model have been well documented in the literature. However, in the multivariate case, only a few papers have dealt with structural breaks. Qu and Perron (2007) consider issues related to estimation, inference, and computation when multiple structural changes occur at unknown dates in a system of equations including a vector autoregressive model. The changes here can pertain to the regression coefficients and/or the covariance matrix of the errors. This section presents a brief description of Qu and Perron's method before proceeding to an application in a VAR model using U.S. macro-data in the next section.

2.1 The Model and Estimator

In the presentation below, we consider a VAR model with three components, $y_t = (y_{1t}, y_{2t}, y_{3t})'$, and one lag, as in our application in section 3. This can be extended to a general VAR model without any major difficulties. The dates and the number of structural changes in the parameters are unknown. We use m to denote the total number of structural changes and T to denote the sample size. The unknown break dates are denoted by the m vector $\Gamma = (T_1, \ldots, T_m)$ and we use the convention that $T_0 = 1$ and $T_{m+1} = T$. Hence there are $m + 1$ unknown sub-periods, $T_{j-1} + 1 \leq t < T_j, j = 1, \ldots, m + 1$. The model can be written as

$$y_t = \pi_{j0} + \pi_{j1} y_{t-1} + \varepsilon_t, \;\; t = T_{j-1} + 1, \ldots, T_j, \tag{1}$$

where $y_t = (y_{1t}, y_{2t}, y_{3t})'$. In each sub-period $T_{j-1} + 1 \leq t < T_j$: $\pi_{j0} = (\pi_{j0}^{(i)})_{i=1,\ldots,3}$ is the vector of constant parameters;

$\pi_{j1} = (\pi_{j1}^{(kl)})_{k=1,\ldots,3;l=1,\ldots,3}$ indicates the 3x3 matrices of the VAR parameters; and ε_t is the vector of the residuals with mean O and covariance matrix denoted by Σ_j. Here we are interested in the estimation of $\Lambda = \{\widehat{m}, \widehat{T}_1, \ldots, \widehat{T}_m, \widehat{\Sigma}_{j=1,\ldots,m+1}\}$. We suppose in this paragraph that m is known, and will discuss later the estimation of m. For the estimation of model (1) it is convenient to rewrite the model as

$$y_t = x_t' \beta_j + \varepsilon_t, \tag{2}$$

where $x_t' = (I_3 \otimes (1, y_{1t-1}, y_{2t-1}, y_{3t-1}))$ and $\beta_j = (\pi_{j0}^{(1)}, \pi_{j1,}^{(11)} \pi_{j1}^{(12)}, \pi_{j1}^{(13)}, \ldots, \pi_{j0}^{(3)}, \pi_{j1,}^{(31)} \pi_{j1}^{(32)}, \pi_{j1}^{(33)})'$. The estimation method we consider is restricted quasi-maximum likelihood. Conditional on the given break dates $\Gamma = (T_1, \ldots, T_m)$, the Gaussian quasi-likelihood ratio is

$$LR_\Gamma = \frac{\Pi_{j=1}^{m+1} \Pi_{t=T_{j-1}+1}^{T_j} f\left(y_t | x_t; \beta_j, \Sigma_j\right)}{\Pi_{j=1}^{m+1} \Pi_{t=T_{j-1}^0+1}^{T_j^0} f\left(y_t | x_t; \beta_j^0, \Sigma_j^0\right)}, \tag{3}$$

where $f\left(y_t | x_t; \beta_j, \Sigma_j\right) = (2\pi)^{-n/2} \left|\Sigma_j\right|^{-1/2} \exp\{-\frac{1}{2}\left[y_t - x_t'\beta_j\right]' \Sigma_j^{-1}\left[y_t - x_t'\beta_j\right]\}$ and $n = 3$ is the number of equations in the VAR model. $\Gamma^0 = (T_1^0, \ldots, T_m^0)$, β_j^0, and Σ_j^0 indicate the true unknown parameters. Qu and Perron's (2007) approach introduces a restriction in $\beta = (\beta_1', \ldots, \beta_{m+1}')'$ and $\Sigma = (\Sigma_1, \ldots, \Sigma_{m+1})$. For example, we obtain a partial structural-change model when the restriction imposes that a particular subset of β_j is the same for all j. We denote by $g\left(\beta, vec(\Sigma)\right) = 0$ the form of the restrictions in β and (or) Σ, where $g(.)$ is an r-dimensional vector and r the number of restrictions. Hence the restricted log-likelihood ratio is given by

$$rlr_\Gamma = \log(LR_\Gamma) + \lambda' g\left(\beta, vec(\Sigma)\right) \tag{4}$$

and the estimates are

$$\{\widehat{T}_1, \ldots, \widehat{T}_m, \widehat{\beta}, \widehat{\Sigma}\} = \arg\ \max_{(T_1,\ldots,T_m,\beta,\Sigma)} rlr_\Gamma. \tag{5}$$

The maximization (5) is taken over all partitions $\Gamma = (T_1, \ldots, T_m)$ such that $|T_j - T_{j-1}| \geq [\varepsilon T]$ and $T_m \leq [T(1-\varepsilon)]$, where ε is an arbitrarily small positive number and [] denotes the integer part of the argument. The parameter ε plays a trimming role and imposes a minimal length for each regime. One important result is that under more general assumptions the estimates of the break dates $\Gamma = (T_1, \ldots, T_m)$ and the coefficients (β, Σ) are asymptotically independent and valid restrictions on the latter do not affect the distribution of the former. Qu and Perron (2007) also discuss a method based on the principle of dynamic programming which searches for the optimal partition $\Gamma = (T_1, \ldots, T_m)$, i.e., the partition that yields the optimal value of the log-likelihood rlr_Γ. They argue that this method is efficient, as it only requires least-squares calculations of order $O(T)$ and matrix inversions of order $O(n)$.

2.2 Selection of the Number of Breaks

To determine the number of breaks m, we can appeal to the likelihood-ratio test of no change versus some specific number of changes, k. As proposed by Qu and Perron (2007), this statistic can be constructed as follows:

$$\sup LR_T \left(k, p_b, n_{bd}, n_{bo}, \varepsilon\right) = 2\left[\log \widehat{L}_T\left(\widehat{T}_1, \ldots, \widehat{T}_k\right) - \log \widetilde{L}_T\right].$$

Here $\log \widehat{L}_T(\widehat{T}_1, \ldots, \widehat{T}_k)$ is the maximum of the log-likelihood obtained with the optimal partition $\{\widehat{T}_1, \ldots, \widehat{T}_k\}$; $\log \widetilde{L}_T$ is the maximum of the log-likelihood under the null hypothesis of no structural change; p_b is the total number of coefficients that are allowed to change (not including the coefficients of the variance-covariance matrix); and n_{bd} and n_{bo} indicate, respectively, the number of parameters that are allowed to change amongst the diagonal and off-diagonal coefficients of the variance-covariance matrix. As noted above, ε is a parameter that allows us to impose a minimal length for each regime. The limiting distribution of $\sup LR_T$ is discussed in detail by the authors and depends on the aforementioned parameters. It is important to note that this testing procedure is particularly flexible. The test can be used with respect to many cases of structural change: (i) changes only in the coefficients of the conditional mean ($n_{bd} = 0, n_{bo} = 0$); (ii) changes only in the coefficients of the covariance matrix of the residuals ($p_b = 0$); and (iii) changes in all of the coefficients ($p_b \neq 0, n_{bd} \neq 0, n_{bo} \neq 0$).

The test for no change versus an unknown number of breaks can also be considered given some upper-bound M for k. These types of tests are called double maximum tests, and the statistic is defined for some fixed weights $W = \{a_1, \ldots, a_M\}$ as

$$D \ \max \ LR_T(M) = \max_{1 \leq k \leq M} [a_k \sup LR_T(k, p_b, n_{bd}, n_{bo}, \varepsilon)].$$

For uniform weights, i.e., $a_i = 1, i = 1, \ldots, M.$, the statistic is always denoted by $UD\max\ LR_T(M)$. The second test is noted by $WD\max$ $LR_T(M)$ and applies weights to the individuals tests such that the marginal p-values are equal across values of m. Bai and Perron (1998) provide more discussion of this class of tests and their critical values.

We can also consider the sequential test based on the null hypothesis of, say, l break dates versus $l+1$ breaks. The general form of this statistic is

$$SEQ_T(l+1|l) = \max_{1 \leq j \leq l+1} \ \sup_{\tau \epsilon \Lambda_{j,\varepsilon}} lr_T\left(\widehat{T}_1, \ldots, \widehat{T}_{j-1}, \tau, \widehat{T}_j, \ldots, \widehat{T}_l\right) - lr_T\left(\widehat{T}_1, \ldots, \widehat{T}_l\right).$$

Here $lr_T(.)$ denotes the log of the likelihood, $\{\widehat{T}_1, \ldots, \widehat{T}_l\}$ denotes the optimal partition if l breaks are supposed, and $\Lambda_{j,\varepsilon} = \{\tau; \widehat{T}_{j-1} + (\widehat{T}_j - \widehat{T}_{j-1})\varepsilon \leq \tau \leq \widehat{T}_j - (\widehat{T}_j - \widehat{T}_{j-1})\varepsilon\}$. In addition, the limiting distribution of the test depends on the number of coefficients that are allowed to change.

In practice, the preferred strategy to determine the number of breaks is first to look at the $UD\max\ LR_T(M)$ or $WD\max\ LR_T(M)$ tests to see if at least one structural break exists. We can then decide the number of breaks based on the $SEQ_T(l+1|l)$ statistic. We select m breaks such that the $SEQ_T(l+1|l)$ tests are insignificant for any $l \geq m$. Bai and Perron (2003) conclude that this method leads to the best results, and they recommend it for empirical applications. However, in the case of a VAR with restrictions on the coefficients, Qu and Perron (2007) consider only the $SupLR_T$ test to obtain the number and the location of the breaks.

2.3 Method of Lag Selection

The presence of structural changes in multivariate models introduces non-linearity. As such, the usual model-selection criteria (Akaike and

Bayesian information criteria, AIC and BIC) are not applicable. One practical result in Qu and Perron (2007) is that the limited distribution of the estimates of the break points is not affected by the imposition of valid restrictions on the parameters. Hence, a non-linear VAR model with a number of lags, say p, can be considered as a valid restriction of a non-linear VAR model with $p + 1$ lags, if the break dates in the two models are significantly unchanged. We use this result to decide on the number of lags in the application section.[5]

3. House Price Macro-Link in the United States: Evidence of Instability

Under this section we apply Qu and Perron's algorithm to the U.S. macro-data. We consider a VAR model with three endogenous variables—real gross domestic product (GDP) or real private consumption (C), real house prices (HP), and the federal funds rate (I)—as a proxy for the monetary policy stance. Closed specifications accounting for the house price macro-relationship have been used by Iacoviello (2005), Goodhart and Hoffman (2007, 2008), Oikarinen (2009), and Calza, Monacelli, and Stracca (2013), among others.

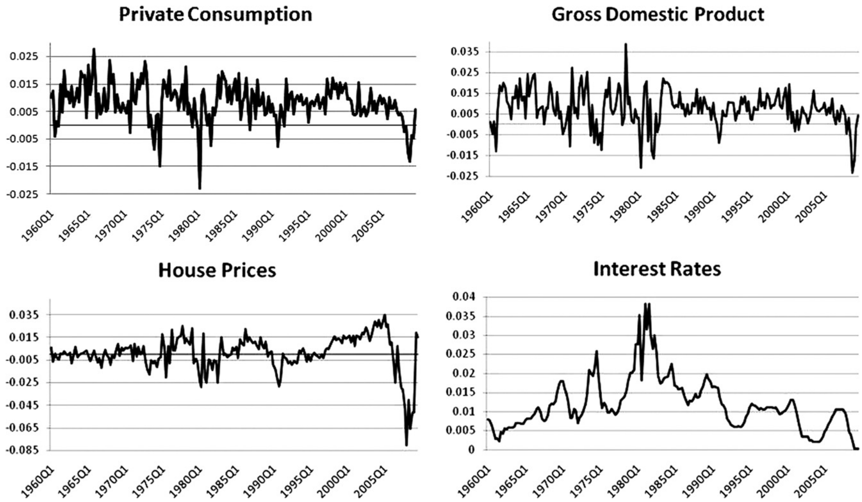
We work with two different VAR specifications, one including GDP and the other private consumption, to examine whether the reduction in credit constraints in the United States could have modified only the impact of housing prices on consumption and not their effect on overall economic activity.

Our data is in quarterly frequency from 1960:Q1 to 2009:Q3. GDP, private consumption, and house prices are measured in logarithms, with the log-values detrended via the first-differencing filter.[6] Private consumption and house prices are deflated using the consumer price index. The federal funds rate is the average value in the

[5] An alternative method of lag selection in a non-linear VAR is proposed in Kurozumi and Tuvaandorj (2011). The authors discuss a modified AIC and BIC to be used in the case of a multivariate model with multiple structural changes.

[6] We use the first-differencing filter instead of a two-sided moving-average filter (for example, Hodrick-Prescott or Baxter-King) in order to test for Granger causality in section 4. Indeed, it is well known in the econometric literature that the use of a two-sided moving-average filter biases the Granger causality test towards a false finding of Granger causation (see, for instance, Cogley 2008).

Figure 1. Data Series



first month of each quarter. The house price series comes from the Case-Shiller Home Price Index. Figure 1 shows the data series used.

3.1 Application of Qu and Perron's Algorithm in a VAR with No Linear Constraints

Qu and Perron's procedure to identify breaks is applied to the following model:

$$y_t = \pi_{j0} + \pi_{j1} y_{t-1} + \varepsilon_t$$
$$\text{and } t = T_{j-1} + 1, \ldots, T_j, j = 1, \ldots, m + 1,$$

where $y_t = (\Delta X_t, \Delta HP_t, \Delta I_t)'$ and m denotes the unknown breaks. In each unknown sub-period $T_{j-1} + 1 \leq t < T_j, \pi_{j0} = (\pi_{j0}^{(i)})_{i=1,\ldots,3}$ is the vector of constant parameters; $\pi_{j1} = (\pi_{j1}^{(kl)})_{k=1,\ldots,3;l=1,\ldots,3}$ indicates the 3x3 matrices of the VAR parameters; and ε_t is the vector of the residuals with mean O and a constant covariance matrix. Thus, we only consider the case of a structural change in the conditional mean of the coefficients. X represents either GDP or private consumption. Δ is the delta operator. All the tests are carried out

Table 1. *WD*max Test for up to Four Breaks: VAR with no Linear Restrictions

	VAR with X = GDP	VAR with X = C
*WD*max Statistic	54.13	50.75
Critical Value at 5%: 30.62		

by fixing $m = 4$ and a trimming value of $\varepsilon = 0.02$.[7] We start by applying the tests considering a lag fixed at $p = 1$.

Let us first consider X = GDP. Table 1 shows the results of the $WD\max\ LR_T(M)$ test, where the null hypothesis is rejected at the 5 percent level. The value of the $WD\max\ LR_T(M)$ statistic is 54.13, whereas the critical value at the 5 percent level is 30.62. Hence, there is at least one structural break in the VAR model. The $SEQ_T\,(l+1|l)$ test allows us to reject the null hypothesis of one break against the alternative of two breaks, as $Seq(2|1) = 40.84$ with the 5 percent critical value being 32.60. The $SEQ_T\,(l+1|l)$ test allows us also to reject the null hypothesis of two breaks against the alternative of three breaks, as $Seq(3|2) = 37.88$ with the 5 percent critical value being 33.96. Finally, the null hypothesis of three breaks against four breaks is not rejected, as $Seq(4|3) = 0.00$ where the 5 percent critical value is 34.89. The performed tests thus lead us to conclude that there are three structural breaks in the considered VAR system which includes a GDP variable: 1969:Q3, 1982:Q4, and 1999:Q1. Table 2 shows the results of the $SEQ_T\,(l+1|l)$ test along with the estimated break dates with their 95 percent confidence intervals.

Next we apply the procedure setting $X = C$. The $WD\max LR_T(M)$ statistic allows us to reject the null hypothesis at the 5 percent level (table 1). The value of the $WD\max\ LR_T(M)$ statistic is 50.75, whereas the critical value at the 5 percent level is 30.62. We conclude then that there is at least one structural

[7]The parameter ϵ imposes a minimal length for each segment, and the limiting distributions of the different tests are affected by this value. As a robustness exercise, we experiment with three different values of ϵ: 0.015, 0.02, and 0.03. These values are considered in many empirical papers and Monte Carlo simulations. We find that results do not differ statistically. Results of this robustness exercise are not presented in the paper but are available by request.

Table 2. The $SEQ_T(l+1|l)$ Test for a VAR with no Linear Restrictions with X = GDP

	Statistic	Critical Value at 5% Level	Significant Break
$Seq(2\|1)$	40.84	32.60	—
$Seq(3\|2)$	37.88	33.96	1969:Q3 (1968:Q1–1971:Q1); 1982:Q4 (1981:Q2–1984:Q2); & 1999:Q1 (1997:Q1–2001:Q1)
$Seq(4\|3)$	0.00	34.89	—

Table 3. The $SEQ_T(l+1|l)$ Test for a VAR with no Linear Restrictions with X = C

	Statistic	Critical Value at 5% Level	Significant Break
$Seq(2\|1)$	39.32	32.60	—
$Seq(3\|2)$	38.82	33.96	1970:Q2 (1967:Q4–1972:Q4); 1985:Q1 (1983:Q4–1986:Q2); & 1999:Q1 (1998:Q2–1999:Q4)
$Seq(4\|3)$	0.000	34.89	—

break in the considered VAR model. The $SEQ_T\ (l+1|l)$ test allows us to reject the null hypothesis of one break against the alternative of two breaks, as $Seq(2|1) = 39.32$ with the 5 percent critical value being 32.60. Moreover, $SEQ_T\ (l+1|l)$ also allows us to reject the null hypothesis of two breaks against the alternative of three breaks, as $Seq(3|2) = 37.88$ with the 5 percent critical value being 33.96. Nevertheless, the null hypothesis of three breaks against four breaks is not rejected, as $Seq(4|3) = 0.00$ where the 5 percent critical value is 34.89. Thus, the performed tests allow us to conclude that there are three structural breaks in the considered VAR specification which includes a private consumption variable: 1970:Q2, 1985:Q1, and 1999:Q1. Table 3 shows the 95 percent confidence intervals for each break.

Afterwards, we apply the tests using a VAR with a lag fixed at $p = 2$. The number and location of the break dates remain significantly unchanged.[8] In both cases, the VAR model with one lag can then be considered as a valid and parsimonious restriction of the VAR model with two lags.

3.2 Application of Qu and Perron's Algorithm in a VAR with Linear Constraints

The estimated break dates in the first part of this section are submitted to a robustness exercise in this sub-section. We would like to dismiss the possibility that the breaks found are not primarily driven by changes in the effect of house prices on the macroeconomy. For instance, the new monetary policy paradigm of the Federal Reserve—nominal policy interest rates reacting more than proportionally to inflation rate changes—could be the main factor causing the rupture detected in the mid-80s. To avoid this issue, we constrain our VAR to only allow a change in the set of parameters related directly to the impact of house prices in the macro-variables. Thus, Qu and Perron's algorithm is applied to the following VAR model:

$$\begin{cases} \Delta X_t = \beta_1 + \beta_2 \Delta X_{t-1} + \beta_{3j} \Delta HP_{t-1} + \beta_4 I_{t-1} + \epsilon_{t1} \\ \Delta HP_t = \beta_5 + \beta_6 \Delta X_{t-1} + \beta_{7j} \Delta HP_{t-1} + \beta_8 I_{t-1} + \epsilon_{t2} \\ \Delta I_t = \beta_9 + \beta_{10} \Delta X_{t-1} + \beta_{11j} \Delta HP_{t-1} + \beta_{12} I_{t-1} + \epsilon_{t3} \end{cases}$$

$$\text{and } t = T_{j-1} + 1, \ldots, T_j, j = 1, \ldots, m+1,$$

where m denotes the unknown breaks and $\epsilon_t = \{\epsilon_{t1}, \epsilon_{t2}, \epsilon_{t3}\}$ is the error term with mean 0 and a constant variance-covariance matrix. Then, in the above VAR with linear restrictions, only the parameters β_{3j}, β_{7j}, and β_{11j} are allowed to change. As proposed by Qu and Perron (2007), the $SupLR_T$ test is performed to obtain the number and the location of breaks in a VAR with linear constraints.

Let us consider the case of a VAR with X = GDP. The null hypothesis of no break is tested against the previously estimated three breaks. Results (table 4) show that the $SupLR_T$ test does not allow us to reject the null hypothesis of no structural break against the alternative of three breaks, as the value of the $SupLR_T$ statistic is 29.92 while the critical value at the 5 percent level is 33.71. Thus,

[8] The test results are not reported because of space considerations but are available by request.

Table 4. The $SupLR_T$ Test for a VAR with no Linear Restrictions with X = GDP

	Statistic	Critical Value at 5% Level	Significant Break
$SupLR_T$ Test (1|0)	25.08	21.37	1984:Q4 (1981:Q4–1987:Q4)
$SupLR_T$ Test (2|0)	29.54	28.17	1971:Q2 (1968:Q2–1973:Q2); 1983:Q1 (1980:Q2–1985:Q4)
$SupLR_T$ Test (3|0)	29.92	33.71	—

we cannot associate each of the three previously detected breaks as structural changes in the effect of house prices on the macroeconomy. We proceed to examine if two of the estimated breaks are robust to the imposed restrictions. The $SupLR_T$ test allows us to reject the null hypothesis of no structural break against the alternative of two breaks, as the value of the $SupLR_T$ statistic is 29.54 while the critical value at the 5 percent level is 28.17. The test gives us 1971:Q2 and 1983:Q1 as structural breaks. The confidence intervals associated with these two breaks, 1968:Q2–1973:Q2 and 1980:Q2–1985:Q4, cover the first two break points obtained in the non-restricted VAR (1969:Q3 and 1982:Q4). We can then conclude that only the break dates 1969:Q3 and 1982:Q4 can be associated with a change in the impact of house prices on the macroeconomy, when GDP is included in the VAR specification. In a final step, we use the $SupLR_T$ test to detect the location of the strongest rupture in the system during the considered period. To do this, we apply the $SupLR_T$ test with the null hypothesis of no structural break against the alternative of one break. The $SupLR_T$ test allows us to reject the null hypothesis of no structural break against the alternative of one break, as the value of the $SupLR$ statistic is 25.08 while the critical value at the 5 percent level is 21.37. The test gives us 1984:Q4 as a break point, and the confidence interval associated with this break, 1981:Q4–1987:Q4, covers the second break date obtained in the non-restricted VAR (1982:Q4). Thus, we conclude that the strongest rupture in the house price macro-link, for the considered VAR specification, occurred in the mid-80s.

We now test the robustness of the three break dates estimated in a VAR including private consumption as an endogenous variable (i.e., X = C). Results (table 5) show that the $SupLR_T$ test does not

Table 5. The $SupLR_T$ Test for a VAR with no Linear Restrictions with $X = C$

	Statistic	Critical Value at 5% Level	Significant Break
$SupLR_T$ Test (1|0)	23.05	21.37	1984:Q3 (1981:Q1–1988:Q1)
$SupLR_T$ Test (2|0)	24.24	28.17	—
$SupLR_T$ Test (3|0)	24.40	33.70	—

allow us to reject the null hypothesis of no structural break against the alternative of three breaks, as the value of the $SupLR_T$ statistic is 24.40 while the critical value at the 5 percent level is 33.70. The $SupLR_T$ test also does not allow us to reject the null hypothesis of no structural break against the alternative of two breaks, as the value of the $SupLR_T$ statistic is 24.24 while the critical value at the 5 percent level is 28.17. Finally, the $SupLR_T$ test allows us to reject the null hypothesis of no structural break against the alternative of one break. The test gives us 1984:Q3 as the only break in the system. The confidence interval associated with this break, 1981:Q1–1988:Q1, covers the second break point obtained in the non-restricted VAR (1985:Q1). We can then conclude that only the break point 1985:Q1 can be associated with a change in the effect of house prices on the macroeconomy when private consumption is included in the VAR system.

Results in this section suggest that the housing prices' impact on the selected U.S. macro-aggregates experienced an important rupture in the mid-80s. Even though the economic agents experienced a progressive relaxation on credit constraints, especially during the last thirty years, the modification of the housing market in the mid-80s seems to have been a determining factor of the change of the house price macro-relationship.[9]

[9] Our results also suggest that the break estimated at the end of the 1960s is not due to a change in house prices' effect on consumption, but it would be mainly driven by changes in the impact of house prices on investment. The increase in the competition on the secondary market due to the creation of Freddie Mac (in 1970) could have played a role in the modification of the link between housing investment and house prices; however, more work would be needed to clarify this issue.

4. Multivariate Time-Series Analysis within Sub-Samples

In section 3 we have shown that the impact of house prices on the macroeconomy changed in the mid-80s. In this section we look for time-series evidence to find in which direction the house prices macro-link has moved. More specifically, we seek evidence that the effect of house prices on economic activity and on private consumption has been reinforced after the break. In this aim, we split the sample into two segments using the break date estimated in the mid-80s. Impulse response and Granger causality tests are performed for both sub-samples. The VAR specification used to identify the breaks in the previous section is also considered in each sub-period.[10]

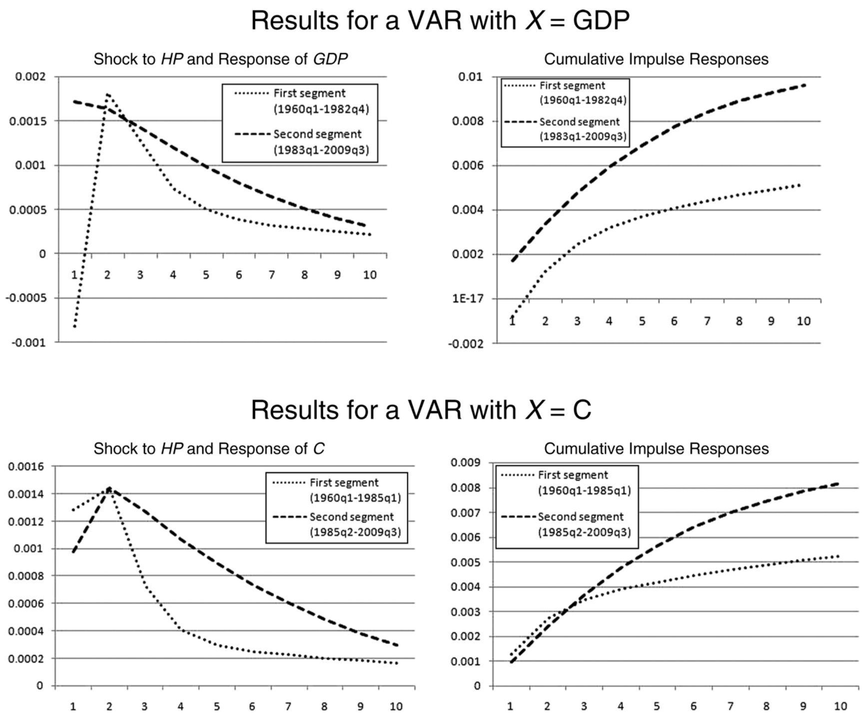
4.1 Generalized Impulse Response Functions

In order to analyze the dynamics of a VAR system, most of the empirical papers in the literature use the "orthogonalized" impulse responses, where the underlying shocks are orthogonalized using the Cholesky decomposition methodology. The main problem with this method is that in practice, when the order of the variables in the VAR is modified, the responses are often significantly different. To avoid the issue of potentially different results, we apply the generalized impulse responses (GIRs) proposed by Pesaran and Shin (1998), which are invariant to the ordering of the variables in the VAR. For each of the VAR specifications, either including GDP or consumption, we perform GIRs and cumulative GIRs for each of the two sub-periods (figure 2). Shocks represent an exogenous increase of 1 percent in the house price variable.[11]

[10] In the previous section we have found that a VAR with one lag is a valid and parsimonious restriction of a VAR with two lags. This result allows us to estimate for this section a VAR with either one or two lags. We choose a VAR with one lag for the sake of preserving an adequate number of degrees of freedom for the estimation.

[11] Because the housing price series has been log-transformed and first-differentiated, an increase of 1 percent of this variable represents an increase of 1 percent in the growth rate of house prices. The same is true for the responses of GDP and private consumption.

Figure 2. Generalized Impulse Responses



Notes: The y-axis measures percent deviations from the baseline, while the x-axis represents quarters after shock.

4.1.1 First Case: GIRs for a VAR Including GDP

First Segment (1960:Q1–1982:Q4). The effect of an increase in house prices by 1 percent has an immediate negative effect on GDP; however, it becomes positive after one quarter and then decreases steadily in the next quarters. The cumulative impulse responses for a ten-quarter horizon suggests that the accumulated effect of a 1 percent increase in housing prices on GDP after ten quarters is around 0.5 percent for this segment.

Second Segment (1983:Q1–2009:Q3). For this sub-period, we find a stronger effect of house prices on GDP compared with the effect found in the first segment for all the periods but one. Moreover, the cumulative impulse responses show that the accumulated effect of a rise of 1 percent in house prices on GDP is around 1 percent after ten quarters for this sub-period, which is approximately 100 percent stronger than the accumulated effect estimated for the first segment.

4.1.2 Second Case: GIRs for a VAR Including Private Consumption

First Segment (1960:Q1–1985:Q1). The effect of a 1 percent increase in the house price variable increases private consumption by approximately 0.1 percent in the first quarter. The shock appears to be absorbed rapidly, implying that a change in housing prices has little influence on private consumption after two quarters. The cumulative impulse responses for a ten-quarter horizon suggests that the accumulated effect of an increase in housing prices of 1 percent on consumption after ten quarters is around 0.5 percent for this sub-period.

Second Segment (1985:Q2–2009:Q3). For this segment, the rise in private consumption following a house price increase of 1 percent is very close in the short run to the response estimated in the first sub-period. Nevertheless, the disturbance appears to impact consumption long after the initial impact. Moreover, the cumulative impulse responses show that the accumulated effect of a rise of 1 percent in house prices on private consumption is around 0.8 percent after ten quarters for this sub-period, which is more than 50 percent stronger than the accumulated effect found for the first segment.

In sum, generalized impulse responses show that the effect of house prices on the economic activity and on aggregate private consumption has substantially increased since the mid-80s.

4.2 Granger Causality Test of Block Exogeneity

Under this short sub-section, multivariate Granger causality of block exogeneity Wald tests are performed for each of the VAR specifications, including either GDP or private consumption, and for each sub-period (see table 6).

Granger causality tests show that before the mid-80s, there is no unidirectional Granger causality running either from house prices to

Table 6. Granger Causality Tests

		Chi-square	P-value
Results for a VAR with X = GDP			
First Segment (1960:Q1–1982:Q4)	$HP \nRightarrow GDP$	0.81	0.37
	$GDP \nRightarrow HP$	0.01	0.93
Second Segment (1983:Q1–2009:Q3)	$HP \Rightarrow GDP$	10.86	0.00
	$GDP \nRightarrow HP$	3.32	0.07
Results for a VAR with X = C			
First Segment (1960:Q1–1985:Q1)	$HP \nRightarrow C$	1.16	0.28
	$C \nRightarrow HP$	1.00	0.31
Second Segment (1985:Q2–2009:Q3)	$HP \Rightarrow C$	20.74	0.00
	$C \nRightarrow HP$	0.66	0.42

Notes: The null hypothesis is the non-Granger causality. The $\Rightarrow$ and $\nRightarrow$ symbols denote the Granger causality and the non-Granger causality, respectively, at the 5 percent significance level.

GDP or consumption, or in the opposite direction. However, a clear unidirectional Granger causality running from house prices to the economic activity and to private consumption is detected since the mid-80s.

5. Conclusion

This paper provides empirical evidence on the strengthening of the impact of house prices on the macroeconomy in the United States. First, Qu and Perron's methodology is applied to identify breaks in a VAR model. The choice of this method is guided by its key advantages compared with the alternative methods used in the literature: the method allows us to detect a break at unknown dates and it allows for multiple breaks. Second, the estimated break points are used to split the sample into different segments, and the generalized impulse responses proposed by Pesaran and Shin (1998) as well as Granger causality tests are performed within the sub-samples.

Using macro-data from the last fifty years, this study finds that there is a robust structural break in the mid-80s in the house price effect on the macroeconomy, which is obtained when either GDP or private consumption is included in the multivariate specification. The paper also finds time-series evidence that the overall impact of house prices on consumption and on economic activity has intensified since the mid-80s. Results in this study support the widespread hypothesis in the theoretical literature that the deep transformation in the mortgage market during the mid-80s is the main cause of the intensification of the impact of housing prices, not only on household expenditures decisions but also on the economic activity as a whole.

Future research could be focused on the consequences of the Dodd-Frank Act on credit conditions and to what extent it will affect the house price macro-link. Indeed, a tightening in credit constraints could be expected following Title XIV of the Dodd-Frank Act—"Mortgage Reform and Anti-Predatory Lending Act"—since its main objective is the imposition of restrictions on mortgage originators to only provide lending to borrowers who are likely to repay their loans.

References

Altissimo, F., E. Georgiou, T. Sastre, M. Valderrama, G. Sterne, M. Stocker, M. Weth, K. Whelan, and A. Willman. 2005. "Wealth and Asset Price Effects on Economic Activity." ECB Occasional Paper No. 29.

Bai, J., and P. Perron. 1998. "Estimating and Testing Linear Models with Multiple Structural Changes." *Econometrica* 66 (1): 47–78.

———. 2003. "Computation and Analysis of Multiple Structural Change Models." *Journal of Applied Econometrics* 18 (1): 1–22.

Bostic, R., S. Gabriel, and G. Painter. 2009. "Housing Wealth, Financial Wealth, and Consumption: New Evidence from Micro Data." *Regional Science and Urban Economics* 39 (1): 79–89.

Calza, A., T. Monacelli, and L. Stracca. 2013. "Housing Finance and Monetary Policy." *Journal of the European Economic Association* 11 (S1): 101–22.

Case, K., R. Shiller, and J. Quigley. 2005. "Comparing Wealth Effects: The Stock Market versus the Housing Market." *Advances in Macroeconomics* 5 (1): 1–32.

Cogley, T. 2008. "Data Filters." In *The New Palgrave Dictionary of Economics,* 2nd ed., ed. S. N. Durlauf and L. E. Blume. Palgrave Macmillan.

Gerardi, K., H. S. Rosen, and P. Willen. 2010. “The Impact of Deregulation and Financial Innovation on Consumers: The Case of the Mortgage Market.” *Journal of Finance* 65 (1): 330–60.

Goodhart, C., and B. Hoffmann. 2007. *House Prices and the Macroeconomy: Implications for Banking and Price Stability.* Oxford: Oxford University Press.

———. 2008. “House Prices, Money, Credit and the Macroeconomy.” *Oxford Review of Economic Policy* 24 (1): 180–205.

Green, R. 1997. “Follow the Leader: How Changes in Residential and Non-Residential Investment Predict Changes in GDP.” *Real Estate Economics* 25 (2): 253–70.

Greenspan, A. 2002. “The Economic Outlook.” Testimony of Federal Reserve Chairman Alan Greenspan before the Joint Economic Committee, U.S. Congress, November 13.

Greenspan, A., and J. Kennedy. 2008. “Sources and Uses of Equity Extracted from Homes.” *Oxford Review of Economic Policy* 24 (1): 120–44.

Haurin, D., and S. Rosenthal. 2006. “House Price Appreciation, Savings, and Consumer Expenditures.” Working Paper, Ohio State University.

Iacoviello, M. 2005. “House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle.” *American Economic Review* 95 (3): 739–64.

Iacoviello, M., and S. Neri. 2010. “Housing Market Spillovers: Evidence from an Estimated DSGE Model.” *American Economic Journal: Macroeconomics* 2 (2): 125–64.

Kiyotaki, N., and J. Moore. 1997. “Credit Cycles.” *Journal of Political Economy* 105 (2): 211–48.

Kurozumi, E., and P. Tuvaandorj. 2011. “Model Selection Criteria in Multivariate Models with Multiple Structural Changes.” *Journal of Econometrics* 164 (2): 218–38.

Miles, D. 1992. “Housing Markets, Consumption and Financial Liberalisation in the Major Economies.” *European Economic Review* 36 (5): 1093–1136.

Miller, N., L. Peng, and M. Sklarz. 2011. “House Prices and Economic Growth.” *Journal of Real Estate Finance and Economics* 42 (2): 522–41.

Muellbauer, J. 2008. “Housing, Credit and Consumer Expenditure.” CEPR Discussion Paper No. 6782.

Oikarinen, E. 2009. "Interaction between Housing Prices and Household Borrowing: The Finnish Case." *Journal of Banking and Finance* 33 (4): 747–56.

Pesaran, H., and Y. Shin. 1998. "Generalized Impulse Response Analysis in Linear Multivariate Models." *Economics Letters* 58 (1): 17–29.

Qu, Z., and P. Perron. 2007. "Estimating and Testing Structural Changes in Multivariate Regressions." *Econometrica* 75 (2): 459–502.

Stock, J., and M. W. Watson. 2003. "Has the Business Cycle Changed and Why?" *NBER Macroeconomics Annual 2002* 17: 159–218.