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House prices, money, credit, and the macroeconomy

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Abstract This paper assesses the links between money, credit, house prices, and economic activity in industrialized countries over the last three decades. The analysis is based on a fixed-effects panel vector autoregression, estimated using quarterly data for 17 industrialized countries spanning the period 1970–2006. The main results of the analysis are the following. (i) There is evidence of a significant multidirectional link between house prices, monetary variables, and the macroeconomy. (ii) The link between house prices and monetary variables is found to be stronger over a more recent sub-sample from 1985 to 2006. (iii) The effects of shocks to money and credit are found to be stronger when house prices are booming.

Key words: house prices, wealth effects, collateral, financial liberalization, money and credit

JEL classification: E21, E22, E31, E32, E44, E47, E50, R21, R31

I. Introduction

Modern-style macro models are inherently non-monetary. Since there are, by construction, no banks, no borrowing constraints, and no risks of default, the risk-free short-term interest rate suffices to model the monetary side of the economy (Goodhart, 2007). As a consequence, money or credit aggregates and asset prices play no role in standard versions of these models. This stands in sharp contrast to the concerns recently expressed by many non-academic observers who have argued that, as a result of the world-wide brisk growth of monetary and credit aggregates over the last couple of years, asset markets are 'awash with liquidity' and that this situation has been responsible for low capital-market yields and inflated asset prices, at least until mid-2007.

In particular, in recent years many industrialized countries have experienced extraordinarily strong rates of money and credit growth accompanied by strong increases in house prices. This

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observation raises a number of questions which are potentially of importance for monetary and regulatory policies. Does the observed coincidence between house prices and monetary variables reflect merely the effects of a common driving force, such as monetary policy or the economic cycle, or does it reflect a direct link between the two variables? If there is a direct link, does it run from house prices to monetary variables, or from monetary variables to house prices, or in both directions? Do fluctuations in house prices and monetary variables have repercussions on the macroeconomy, i.e. for the development of real GDP and consumer prices? And, finally, what is the relevant monetary variable in this context, money, or credit, or both?

From a theoretical point of view, the interlinkages between monetary variables, house prices, and the macroeconomy are multi-faceted. Optimal portfolio adjustment mechanisms, which are at the heart of the traditional monetarist view of the transmission process, suggest a two-way link between house prices and money. An expansion of money changes the stock, and the marginal utility of liquid assets relative to the stock, and the marginal utility of other assets. Agents attempt to restore equilibrium by means of adjustments in spending and asset portfolios that re-equate for all assets as well as for consumption the ratios of marginal utilities to relative prices. This implies that an increase in money triggers increases in a broad range of asset prices and decreases in a broad range of interest rates and yields. In this sense, monetarists characterize the development of money as reflecting changes in the whole spectrum of interest rates and asset prices which are relevant for spending and investment decisions (Meltzer, 1995; Nelson, 2003). By the same token, a change in house prices alters the value of the stock of housing assets, triggering a portfolio rebalancing which will also involve an adjustment in the demand for monetary assets (Greiber and Setzer, 2007).[1]

A link between credit and house prices may arise via housing wealth and collateral effects on credit demand and credit supply, and via repercussions of credit supply fluctuations on house prices. According to the life-cycle model of household consumption, a permanent increase in housing wealth leads to an increase in household spending and borrowing when homeowners try to smooth consumption over the life cycle. Besides this wealth effect, there is also a collateral effect of house prices emanating from the fact that houses are commonly used as collateral for loans because they are immobile and can, therefore, not easily be put out of a creditor's reach. As a consequence, higher house prices not only induce homeowners to spend and borrow more, but also enable them to do so by enhancing their borrowing capacity.[2,3]

It is, however, often claimed that, while an increase in the physical stock of houses undeniably represents an augmentation of the nation's wealth, the effects of a change in housing wealth induced by a change in house prices is not clear a priori. This is because a permanent increase in house prices will not only have a positive wealth and collateral effect on landlords and owner-occupiers, but it will also have a negative income effect on tenants who have to pay higher rents, and on prospective first-time buyers who now have to save more for their

[1] Besides the housing wealth effect on money demand, there might further be a transactions effect, arising generally from higher demand for transactions balances when wealth increases, and specifically from higher transactions related to house purchase when house prices rise.

[2] Aoki *et al.* (2004) and Iacoviello (2004, 2005) show, based on general equilibrium models, that a financial accelerator effect arises in the household sector via house prices, when households' ability to borrow depends on the value of housing collateral.

[3] For a more detailed exposition of the wealth and collateral effect of house prices on consumption, see Muellbauer (2007).

intended house purchase. Thus, those who have already satisfied their housing requirements gain, while those who have yet to do so, or who are renting, lose. But then it is not clear why the same analysis is not applied when the relative prices of equities, or bonds, rise, in so far as both housing and financial asset prices increase because of a fall in discount rates. Also it might be argued that those who have already completed their life-cycle purchases gain from asset-price increases, while those who have yet to save up for retirement lose. Moreover, a large proportion of the 'losers' from a relative housing-price increase are those yet to be born, and those too young to be earning for themselves. They can hardly save more, or lower their current consumption, whereas the old homeowners (the net gainers) can, and will, raise their consumption. There is, therefore, an asymmetry between gainers and losers, which works in favour of a positive wealth or collateral effect of house prices on consumption.

While the housing wealth and housing collateral effects on consumption are the most important or most explored channels of the transmission of house-price fluctuations to the real economy, the transmission via private investment also plays a role. The most direct effect of house-price fluctuations on economic activity is via residential investment. An increase in house prices raises the value of housing relative to construction costs, i.e. the Tobin q for residential investment. New housing construction becomes profitable when house prices rise above construction costs. Residential investment is therefore a positive function of house prices. Furthermore, the value of collateralizable property and land also affects the ability of firms to borrow and finance business investment, giving rise to a positive link between house prices and business investment.[4]

These wealth and collateral effects of house prices on consumption and investment imply adjustments in credit demand and credit supply, thereby potentially giving rise to a causal link from house prices to credit aggregates. House prices influence credit demand via wealth effects on consumption and Tobin's q effects on investment, while the collateral effects also have an impact on credit supply. Additional credit supply effects may arise via the effect of house prices on the balance sheets of banks. Such an effect may result directly via banks' property wealth, and indirectly via the effect on the value of loans secured by real estate.[5]

An exogenous change in credit supply, e.g. driven by financial liberalization, may in turn also have repercussions on house prices. The price of property can be seen as an asset price, which is determined by the discounted future stream of property returns. An increase in credit supply lowers lending interest rates and stimulates current and future expected economic activity. As a result, property prices may rise because of higher expected returns on property and a lower discount factor. An increase in the availability of credit may also increase the demand for housing if households are borrowing-constrained. With supply temporarily fixed because of the time it takes to construct new housing units, this increase in demand will be reflected in higher property prices.

These (very basic) theoretical considerations suggest that there are probably good reasons to believe that there exists a multidirectional link between money, credit, house prices, and

[4] Bernanke and Gertler (1989), Kiyotaki and Moore (1997), and Bernanke *et al*. (1999) have developed modified real business cycle models wherein firms' borrowing capacity depends on their collateralizable net worth, and show that fluctuations in firms' net worth amplify macroeconomic shocks and can give rise to a powerful financial accelerator effect.

[5] For example, Chen (2001) develops a general equilibrium model in which both borrowers' and banks' net worth influences the supply of credit. Just as borrowers' net worth acts as an incentive mechanism and collateral for the banks, bank capital acts in these models as an incentive mechanism and as collateral for the bank's providers of loanable funds, e.g. depositors. So, the availability of loanable funds to banks depends on their capitalization.

the wider economy. However, while these theoretical considerations give us some tentative indications, they obviously do not allow any definite conclusions. In the absence of a fully fledged theoretical model integrating all the potential interlinkages between house prices, money, credit, and the macroeconomy we have described above, the issue ultimately has to be addressed empirically.

There are already quite a number of empirical studies on this subject. However, none of the existing studies addresses all the relevant questions we have raised above. Most studies focus on the link between credit and property prices, but explore the link only in one direction.[6] Other studies investigate the two-way character of the link, but do not address the question of whether credit or money is the relevant monetary variable.[7] Finally, there are a number of studies addressing the latter, but without addressing the potential two-way character of any potential link between house prices and monetary variables.[8]

In this paper we try to contribute to closing this gap by assessing the link between money, credit, house prices, and the economy in a multivariate context. The analysis is performed for a panel of 17 industrialized countries based on a fixed-effects panel vector autoregression (VAR) estimated over the sample period 1973–2006, which is the longest sample period for which the variables required for the analysis are available and which is also the period covered by many of the above-referenced studies. We further re-estimate the model over a shorter sub-sample, 1985–2006, and compare the results with those obtained for the full sample period. This is done because there are good reasons to believe that the link between monetary variables and house prices, and also macroeconomic dynamics in general, have changed in the late 1970s and early 1980s. On the one hand, there has been a major change in the paradigm governing the conduct of monetary policy since the late 1970s. After the experiences of the 'Great Inflation' in the 1970s, with high and volatile inflation accompanied by high volatility of output and unemployment, restoring price stability became the overarching goal

[6] Borio *et al*. (1994) investigate the relationship between credit-to-GDP ratios and aggregate asset prices for a large sample of industrialized countries. They find that adding the credit-to-GDP ratio to an asset pricing equation helps to improve the fit of this equation in most countries. Based on simulations, they demonstrate that the boom–bust cycle in asset markets of the late 1980s to the early 1990s would have been much less pronounced or would not have occurred at all had credit ratios remained constant. Goodhart (1995) investigates the effect of property prices on bank lending in the UK and the USA using long spans of historical data, and finds that property prices significantly affect credit growth in the UK but not in the USA. Hofmann (2004) analyses the role of property prices in explaining credit dynamics in industrialized countries since 1980. He finds that property prices are an important determinant of the long-term trend development in credit over this period and that increases in property prices have a highly significant positive effect on credit dynamics.

[7] Hofmann (2003), Goodhart and Hofmann (2004*a*), and Goodhart *et al*. (2006) analyse the relationship between bank lending and property prices based on a multivariate empirical framework and find that causality does, in fact, seem to go in both directions, but that the effect of property prices on credit appears to be stronger than the effect of credit on property prices. Gerlach and Peng (2005) analyse the link between property prices and credit in Hong Kong and find that causality runs from property prices to lending, rather than conversely. Greiber and Setzer (2007) investigate the link between broad money and property prices in the USA and the euro area. They find that adding property prices to an otherwise standard money demand system restores a stable money demand equation in both economies. Based on a standard impulse–response analysis, they further show that causality runs in both directions: an increase in broad money growth triggers an increase in property prices and vice versa.

[8] Gouteron and Szpiro (2005) investigate the effect of excess liquidity, measured by the ratio of broad money to GDP and, alternatively, the ratio of private credit to GDP, in the USA, the euro area, the UK, and Japan, but fail to detect any significant links except for the UK. Adalid and Detken (2007) explore the effect of broad money growth on house prices in a panel of industrialized countries and find that the link is significant and particularly strong in times of aggregate asset-price booms. They further find that private credit growth does not have a significant effect on house-price dynamics.

of monetary policy in industrialized countries. This change in paradigm was reflected in a significant decline in inflation rates around the world in the early 1980s. Over the same period, there has been a substantial reduction in macroeconomic volatility in many countries, a phenomenon referred to as the 'Great Moderation', which is probably, at least in part, also attributable to the change in the monetary policy paradigm. On the other hand, as we have argued in Goodhart *et al*. (2004), financial systems have undergone substantial changes over the last decades since the 1970s. Financial systems in the industrialized countries have been liberalized and deregulated, which could have strengthened the link between property prices and the financial sector.[9]

In Goodhart *et al*. (2004) we have further argued that financial-sector liberalization is likely to have increased the procyclicality of financial systems by fostering procyclical lending practices of banks. In fact, the historical experience has shown that financial imbalances and asset-price booms or bubbles have often been preceded by brisk expansions of credit and money.[10] Against this background, some commentators have recently argued that a monetary-policy strategy that attaches some weight to the monitoring of monetary variables, rather than following a pure inflation-targeting approach, may help to avoid adverse longer-run consequences of building up financial imbalances by automatically inducing a leaning-against-the-wind monetary policy mitigating excessive asset-price bubbles.[11]

We test the hypothesis that monetary shocks have stronger effects on house prices in times of house-price booms based on a dummy variable augmented panel VAR. Whether the effects of credit and money shocks are stronger during house-prices booms is then assessed by comparing the impulse responses obtained under the boom scenario with those under the no-boom scenario. This is done based on a dummy variable capturing mechanically identified house-price boom episodes across countries, and based on a dummy variable capturing cross-country differences in average house-price inflation over the period by singling out those countries that have experienced particularly high rates of house-price increases.

We are aware of the fact that, with the chosen empirical set-up, we are obviously not able to disentangle the different structural links we have discussed above, nor are we able to disentangle these structural links from non-structural links arising from forward-looking behaviour. For example, a positive effect of house prices on GDP may reflect the housing wealth and collateral channels described above, but it may also be due to forward-looking agents in the housing market anticipating future movements in GDP and the repercussions of such movements on the future returns of housing assets. We do not see this as a major problem, since the declared aim of this paper is to uncover the lead–lag relationships between money, credit, house prices, and key macroeconomic variables, and to detect changes and non-linearities in these relationships. We do not aim to disentangle the different channels potentially driving any of the estimated statistical associations, which would certainly be a much more challenging, if not an impossible task.

The remainder of the paper is structured as follows. Section II describes the data set and discusses data issues. Section III describes the empirical methodology. In section IV we

[9] In a similar vein, Muellbauer and Murphy (1989), and more recently Muellbauer (2007), have argued that the housing collateral effect on consumption will be stronger when credit markets are liberalized.

[10] See Borio and Lowe (2004), Detken and Smets (2004), and Adalid and Detken (2007).

[11] For example, *The Economist* (2006) recently stated that '(t)his link between money and asset prices is why the ECB's twin-pillar framework may be one of the best ways for central banks to deal with asset prices'. See also Mayer (2005), who characterizes the ECB's strategy as providing 'a bridge between inflation targeting and a new paradigm which takes account of financial and asset market developments in monetary policy decisions'.

present the empirical results. Section V analyses whether the link between monetary variables and house prices is stronger in times of house-price booms. Section VI concludes.

II. Data

The empirical analysis is based on quarterly data for the following 17 industrialized countries: the USA, Japan, Germany, France, Italy, the UK, Canada, Switzerland, Sweden, Norway, Finland, Denmark, Spain, the Netherlands, Belgium, Ireland, and Australia, spanning the period 1970Q1–2006Q4. The set of data series used in the empirical analysis comprises real GDP, the consumer price index (CPI),[12] a short-term nominal interest rate, nominal house prices, nominal broad money, and nominal bank credit to the private sector. Except for the short-term interest rate, all data are seasonally adjusted.

Real GDP data for the euro-area countries were taken from the Eurostat database, back-dated using real GDP series taken from the OECD Quarterly National Accounts (QNA) database (France prior to 1978Q1 and Spain prior to 1980Q1) and the Bank for International Settlements (BIS) Macrodatabase (Finland prior to 1975Q1), if necessary. In some cases (Belgium prior to 1980, the Netherlands prior to 1977, and Ireland prior to 1997), we had to construct quarterly GDP data by interpolation from annual GDP data (taken from the BIS database) based on the Chow–Lin procedure using industrial production (taken from the IMF International Financial Statistics (IFS) database) as the reference series. The real GDP series for the other countries were collected from the BIS Macrodatabase (Sweden), the OECD QNA database (Australia, Canada, Japan, UK), and the St Louis FRED database (USA). For Denmark, we linked a quarterly series from the OECD QNA starting in 1977Q1 to an interpolated quarterly series derived from annual GDP data and quarterly industrial production data (both taken from the IMF IFS database) using the Chow–Lin procedure. For Norway, we linked a quarterly series from the OECD QNA database to a quarterly series taken from the Global Financial Data (GFD) database in 1978Q1.

The CPI data series for the euro-area countries were taken from the ECB database. The ECB series for Germany was linked to a CPI series from the BIS database in 1980Q1, while for Ireland the ECB series was linked to a series taken from the OECD Main Economic Indicators (MEI) database in 1976Q1. For all other countries, the CPI data were taken from the OECD MEI, except for the USA where the source was the St Louis FRED Database.

The short-term interest rate is for most countries a 3-month Treasury Bill (T-Bill) rate taken from the GFD database.[13] When a T-Bill rate was not available, we complemented the database using money-market rates. Three-month money-market rates were used for Finland (from the OECD Economic Outlook database 1970Q1–1998Q4; from the GFD database thereafter) and Spain prior to 1982Q3 (from the GFD database). Overnight money-market rates were used for Switzerland (from the GFD Database), Denmark prior to 1975Q4 (from the IMF IFS database), and Norway prior to 1983Q4 (from the GFD database).

The house-price data were taken from the European Central Bank (ECB) and the BIS databases and from national sources. It is important to note that the available data for house

[12] We used the CPI rather than alternative measures of the aggregate price level, such as the GDP deflator or the consumption deflator, mainly for the reason that central banks' inflation targets or objectives usually refer to some kind of consumer price index. A drawback of using the CPI is that there are occasional changes in methodology, for example in the USA with regard to the measurement of home-ownership costs in 1983.

[13] A short-term money-market rate was for most countries not available for the full sample period.

prices are in most cases not directly comparable across countries, owing to differences in the definition of the representative property, but also due to differences in data collection (Arthur, 2003). In most countries, different house-price series had to be linked in order to obtain a series for the full sample. Also, in some cases no quarterly data were available, so that interpolated semi-annual data (Italy, Japan) or interpolated annual data (Germany and Belgium pre-1980, France pre-1981, Spain pre-1987, Ireland pre-1975) had to be used. The interpolation was performed based on the Chow–Lin procedure using either a construction cost index, or the CPI sub-index for rent, or both when available as reference series. More details on the house-price series are provided in Appendix Table A1.

Figure 1 shows the year-on-year percentage change in nominal house prices (solid lines) and the year-on-year percentage change in real house prices (dotted lines), calculated by deflating nominal house prices with the CPI. The figures reveal that, while there were a couple of occasions when real house prices declined, a nominal house-price deflation is an extremely rare event and has always been associated with episodes of severe economic downturns or crises, such as the recessions in the early 1980s and early 1990s, and the Nordic and the Japanese banking crises in the 1990s. The USA never experienced a nominal house-price deflation over this sample period.

The data series for credit from the banking sector to the private non-financial sector (private credit) were taken from the IMF IFS (banking institutions' claims on the private sector, series code 22D). The exception is the UK, where we used a series for banks' and building societies' lending to the private sector, taken from the BIS database. Many of the IMF credit series displayed large level shifts owing to changes in definitions or re-classifications.[14] Following Stock and Watson (2003), we adjusted for these level shifts by replacing the quarterly growth rate in the period when the shift occurred with the median of the growth rate of the two periods prior to and after the level shift. The level of the series was then adjusted by backdating the series based on the adjusted growth rates.[15] In a few countries there were also some observations missing, and these were generated based on comparable credit series taken from other sources.[16]

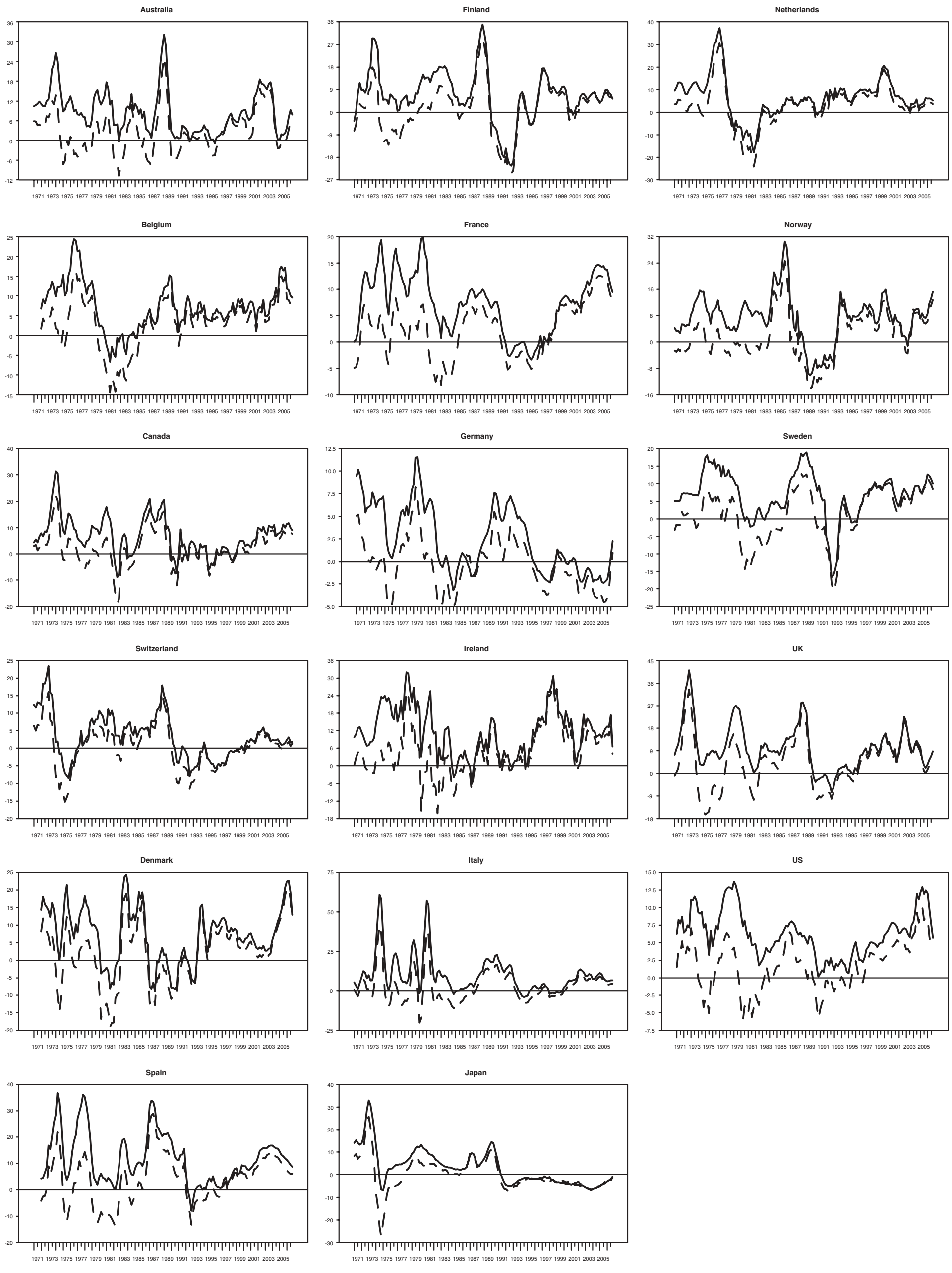
Data series for the broad monetary aggregate, M3, in the EMU member countries were taken from the ECB database. The German series was adjusted for a level shift in 1990Q3 using the Stock–Watson methodology described in the previous paragraph. For the other countries, seasonally adjusted data series for the most relevant broad monetary aggregate were taken

[14] For the euro-area countries, there is a credit series in national currency until 1998Q4 and a series in euros from 1999Q1. Even after converting the national currency series to euros based on the irrevocable fixed exchange rates, some of the spliced credit series still displayed level shifts in 1999Q1. For this reason we performed a level-shift adjustment in this quarter for all euro-area countries.

[15] The following level shifts were adjusted for: Australia 1989Q1, 2002Q1; Belgium 1992Q4, 1999Q1; Canada 2001Q4; Denmark 1987Q4, 1991Q1, 2000Q3; Finland 1999Q1; France 1978Q1, 1999Q1; Germany 1990Q2, 1999Q1; Italy 1999Q1; Ireland 1982Q4, 1995Q1, 1999Q1; Japan 1997Q4, 2001Q4; Netherlands 1982Q4, 1999Q1; Norway 1976Q1; Spain 1983Q1, 1986Q1, 1999Q1; Sweden 1983Q1, 1996Q1; Switzerland 1974Q4, 1982Q3, 1996Q4; USA 2001Q4.

[16] Belgium: a missing observation for 1998Q4 was generated using the growth rate of a series for bank lending to the private sector taken from the BIS database. France: a missing observation for 1977Q4 was generated based on the growth rate of a series from the BIS Database named 'Credit of a banking character to the economy'. Netherlands: missing observations for 1998Q1–1998Q4 were generated with the growth rates of a series for claims of monetary institutions on the private sector taken from the BIS Database. Norway: missing observations in 1987Q1–Q2 were generated from the growth rate of an IMF series for credit extended by non-bank financial institutions to the private sector (IMF IFS series code 42D). Sweden: missing observations in 2001Q1–Q3 were generated from the growth rate of a series for bank lending to the private sector from the Riksbank's website.

Figure 1: House-price inflation in industrialized countries, 1971–2006



Note: The graphs display the year-on-year percentage changes in nominal house prices (solid lines) and in real house prices (dotted lines).

from the OECD MEI database (Australia, M3; Canada, M3; Denmark, M3 adjusted for a level shift in 1989Q1 based on the Stock–Watson method described above; Japan, M2 + Certificates of Deposit; Norway, M2; Switzerland, M3; UK, M4) and the St Louis FRED database (USA, M2M).

Figure 2 displays the year-on-year percentage change in nominal broad money (solid lines) and nominal bank credit to the private sector (dotted lines). The main message to be taken from these graphs is that while there is correlation between the two series, this correlation is far from perfect, which means that including both variables in an empirical model, as we do in the following sections, will not give rise to major multicollinearity problems.[17]

III. Methodology

The analysis is based on a panel VAR given by:

$$Y_{i,t} = A_i + A(L)Y_{i,t} + \varepsilon_{i,t}, \tag{1}$$

where $Y_{i,t}$ is a vector of endogenous variables and $\varepsilon_{i,t}$ is a vector of errors. A_i is a matrix of country-specific fixed effects, $A(L)$ is a matrix polynomial in the lag operator whose order is determined by the Akaike information criterion considering orders up to four. The vector of endogenous variables comprises the log difference of real GDP (Δy), the log difference of the consumer price index (Δcpi), the level of the short-term nominal interest rate (ir), the log difference of nominal residential house prices (Δhp), the log difference of nominal broad money (Δm), and the log difference of nominal private credit (Δc). The vector $Y_{i,t}$ is therefore given by

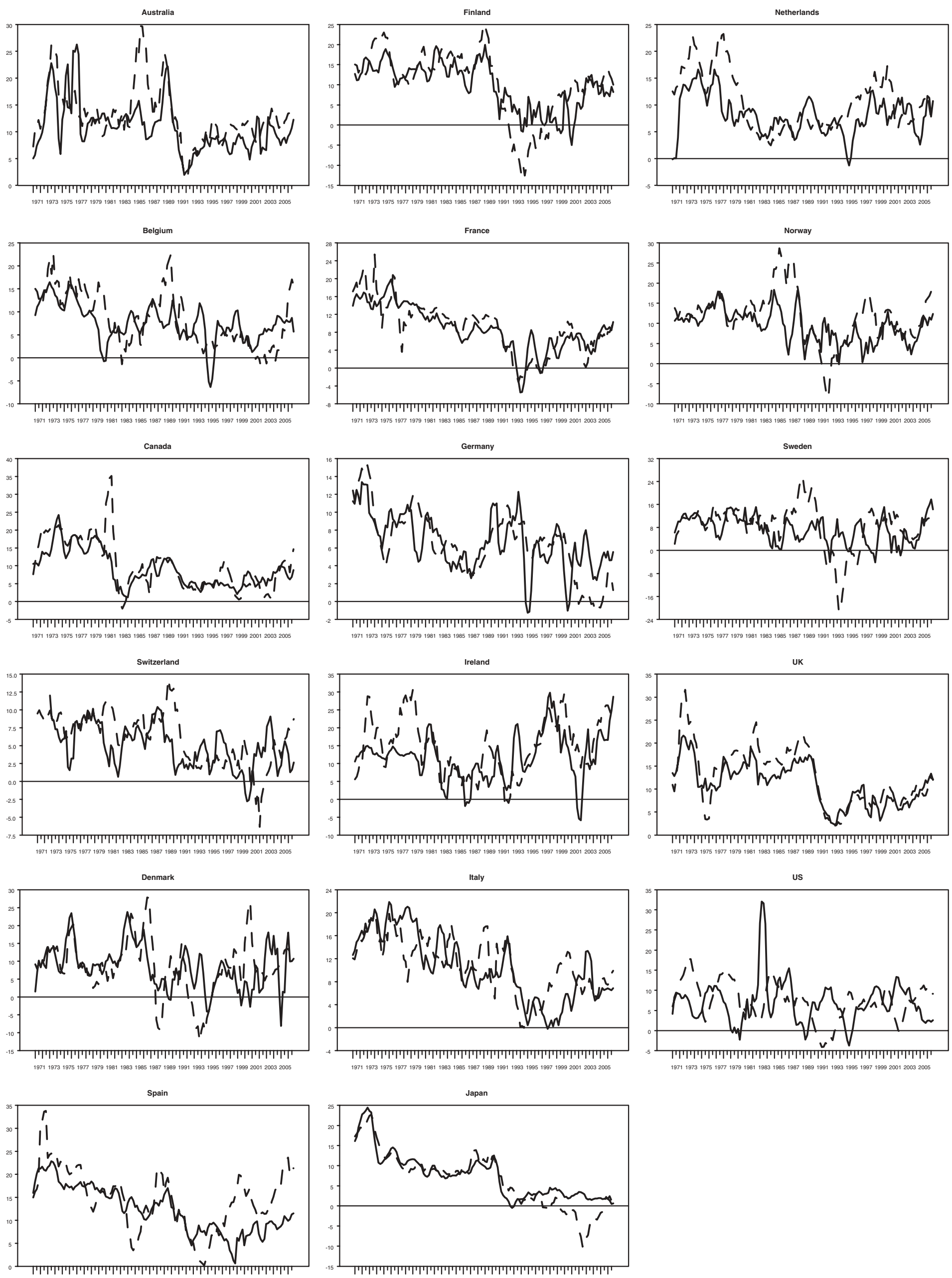
$$Y = [\Delta y, \Delta cpi, ir, \Delta hp, \Delta m, \Delta c]'. \tag{2}$$

The advantage of using a panel-modelling framework is that it substantially increases the efficiency and the power of the analysis. Estimating the six dimensional VAR at the individual country level would suffer from too few degrees of freedom, in particular when the models are re-estimated over a shorter sub-sample starting in 1985. A drawback of the panel approach is that it imposes pooling restrictions across countries and thereby disregards cross-country differences in the estimated dynamic relationships. Indeed, when we checked the validity of the pooling restrictions implied by the panel set-up, we found that they were consistently rejected. However, when we performed the analysis at the individual country level, we also found that the estimated dynamic relationships were often insignificant and in some cases even implausible.[18] In contrast to this, as we show in the following sections, the results from the panel analysis uncovered highly significant dynamic interactions and the results generally made good sense. But rather than interpreting these findings as invalidating the panel set-up, we tend to share the view of Gavin and Theodorou (2005), who argue that adopting a panel approach in a macro framework such as our own helps to uncover common dynamic

[17] For the panel of 17 countries, the cross-correlation of the growth rates of nominal broad money and nominal bank credit is 0.56 for the year-on-year growth rates and 0.38 for the quarterly growth rates.

[18] The results of the individual country analysis are available upon request.

Figure 2: Money and credit growth in industrialized countries, 1971–2006



Note: The graphs display the year-on-year percentage changes in broad money (solid lines) and in private credit (dotted lines).

relationships which might otherwise be obscured by idiosyncratic effects at the individual country level.[19]

We estimate model (1) by OLS. It is well known that OLS estimation of VAR models is subject to Hurwicz-type bias inherent in dynamic models. Since the Hurwicz-type bias goes to zero as the number of observations approaches infinity, it usually does not receive much attention in time-series VAR studies as the number of observations is perceived to be sufficiently large. Unsurprisingly, there has been considerable attention paid to this issue in panel econometrics, since panel data are usually characterized by a large number of cross-sections and a small number of time-series observations. Nickell (1981) has derived analytical expressions for the size of the bias for the first-order autoregressive case, concluding that it is large for panels with small time-series dimensions, even when the number of cross-sections goes to infinity. His results also show that the size of the bias depends negatively on the size of the time-series dimension.

In order to overcome this drawback of the fixed-effects (FE) OLS panel estimator, a number of alternative estimators have been proposed in the literature which are based on instrumental variable estimation (Anderson and Hsiao, 1981) or GMM (Arellano and Bond, 1991; Arellano and Bover, 1995).[20] We nevertheless decided to continue to use the FE OLS estimator for several reasons. In the present application the time-series dimension of the panel is large, with 136 observations per cross-section unit for the full sample estimation starting in 1973Q1 and 88 observations for the sub-sample estimation starting in 1985Q1. Since the size of the fixed-effects bias depends negatively on the number of time-series observations in the panel (Nickell, 1981), it is likely to be of limited importance in our application. Moreover, while the above-mentioned instrumental variable or GMM-based estimators can overcome the bias of the FE estimator, they are in turn subject to other drawbacks. First, instrumental variable estimators are less efficient than OLS estimators, i.e. they tend to produce estimates with a larger variance. This drawback can outweigh the bias of the FE estimator in empirical applications when the time dimension of the model is not too small. Judson and Owen (1999) compare the performance of the FE estimator with the Anderson–Hsiao and the Arellano–Bond estimators in terms of bias and root-mean-squared error (RMSE) of the coefficient estimates based on a Monte-Carlo experiment. They conclude that, even for moderate time-series sample sizes of 30, the FE estimator performs just as well or better than the alternative instrument-based estimators. Second, when the instruments used in the instrumental variables or GMM estimation are only weakly correlated with the instrumented variables, this can in turn give rise to biased coefficient estimates and hypothesis tests with large size distortions (Stock and Yogo, 2002).

Another issue we had to consider was whether we should include time dummies in (2) or not. In typical panel-data studies, when the cross-section dimension is large and the time dimension is small, time dummies are automatically included. This is done because it involves only a minor loss in efficiency since only a small number of dummies have to be added to the model, and because the relationships to be uncovered by the analysis are driven

[19] Gavin and Theodorou (2005) find in their empirical application that, while the pooling restrictions are generally rejected in-sample, the panel model performs significantly better than the individual country model in out-of-sample forecasting.

[20] As an alternative to instrument-based estimators, Kiviet (1995) has proposed a bias-corrected FE estimator which is based on a two-step procedure making use of the analytical bias expressions derived by Nickell (1981). Because of the problems associated with the practical implementation of this approach it is, however, almost never used in empirical applications.

by the cross-section dimension. In our case, the time dimension is large, which means that including time dummies would involve a considerable loss in efficiency. Furthermore, the interlinkages we wish to investigate are in part driven by developments shared at least by sub-groups of countries, so that dummying out common time effects may substantially reduce the information content of the dataset. To be on the safe side, we replicated all the exercises we report in the following sections with a full set of time dummies included, and found that the results were qualitatively not altered.[21]

The results reported in the following section are therefore based on a panel VAR estimated by fixed-effects OLS without time dummies. Based on the thus estimated panel VAR, we first perform standard Granger causality tests. A variable x is said to 'Granger cause' another variable y if the hypothesis that the coefficients on the lags of variable x in the VAR equation of variable y are all equal to zero (i.e. that the lags of variable x can be excluded from the VAR equation of variable y) is rejected by a Wald test. In order to take into account potential heteroskedasticity of the VAR residuals over time and across countries, the tests are based on heteroskedasticity-robust variance–covariance matrices.

Based on Granger causality tests, we can assess the significance of the direct lead–lag relationships between the endogenous variables. These tests, however, do not take into account the indirect effects running via the other variables included in the system, and also do not provide any information about the direction and the strength of the effects. In order to get a more complete picture of the dynamic interactions, we perform, as the next step, an impulse–response analysis based on the estimated VARs. We recover the orthogonalized shocks of the systems based on a simple Cholesky decomposition, with the ordering as given in (2). The ordering of the first three variables is standard from the monetary transmission literature. The ordering of house prices, money, and credit is, of course, somewhat arbitrary and was based on the consideration that the price of a house is probably stickier than monetary variables. Credit was ordered last because it appeared more plausible to allow for an immediate effect of a change in the money stock on credit rather than vice versa. Robustness checks suggested, however, that changes in the ordering of the variables had no substantial effect on the results. Finally, since we want to compare the transmission of the shocks for two different sample periods, we simulate one-unit shocks rather than one-standard-deviation shocks.

The orthogonalized shocks should not be interpreted as structural shocks, but rather as orthogonalized reduced-form shocks. For example, the money shock should be interpreted as an increase in broad money which is unrelated to changes in GDP, goods prices, house prices, and interest rates. It is not possible to disentangle whether the underlying structural driving force is a money-demand or a money-supply shock. Identification of structural shocks might, in principle, be possible, when based on a smaller model set-up and a different shock-identification scheme, such as a combination of long-run and short-run restrictions or sign restrictions. But the use of these more sophisticated identification schemes would require estimating a smaller model, which would drive us away from our original goal, which was to uncover the dynamic lead–lag link between house prices, money credit, and the three key macroeconomic variables. Furthermore, even based on a more sophisticated shock–identification scheme, it would prove difficult to identify all relevant structural shocks, i.e. to disentangle aggregate demand and supply shocks, a monetary policy shock, money demand and supply shocks, a credit demand and supply shock, and a housing demand and supply shock.

[21] The results of the fixed-effects panel VAR estimation with a full set of time dummies included are also available upon request.

Table 1: Granger causality tests (1973–2006 sample)

$\Delta y \rightarrow \Delta cpi$ **3.35** **(0.01)**	$\Delta y \rightarrow ir$ **8.14** **(0.00)**	$\Delta y \rightarrow \Delta hp$ **6.07** **(0.00)**	$\Delta y \rightarrow \Delta m$ 0.44 (0.77)	$\Delta y \rightarrow \Delta c$ **6.94** **(0.00)**
$\Delta cpi \rightarrow \Delta y$ **4.80** **(0.00)**	$\Delta cpi \rightarrow ir$ 7.63 (0.26)	$\Delta cpi \rightarrow \Delta hp$ **4.36** **(0.00)**	$\Delta cpi \rightarrow \Delta m$ **11.39** **(0.00)**	$\Delta cpi \rightarrow \Delta c$ **3.49** **(0.01)**
$ir \rightarrow \Delta y$ **7.80** **(0.00)**	$ir \rightarrow \Delta cpi$ **3.23** **(0.01)**	$ir \rightarrow \Delta hp$ **10.51** **(0.00)**	$ir \rightarrow \Delta m$ 0.89 (0.46)	$ir \rightarrow \Delta c$ **4.40** **(0.00)**
$\Delta hp \rightarrow \Delta y$ **6.23** **(0.00)**	$\Delta hp \rightarrow \Delta cpi$ 1.32 (0.26)	$\Delta hp \rightarrow ir$ 1.73 (0.14)	$\Delta hp \rightarrow \Delta m$ **4.99** **(0.00)**	$\Delta hp \rightarrow \Delta c$ **8.86** **(0.00)**
$\Delta m \rightarrow \Delta y$ **6.70** **(0.00)**	$\Delta m \rightarrow \Delta cpi$ **4.46** **(0.00)**	$\Delta m \rightarrow ir$ 1.40 (0.23)	$\Delta m \rightarrow \Delta hp$ **2.57** **(0.04)**	$\Delta m \rightarrow \Delta c$ **3.02** **(0.02)**
$\Delta c \rightarrow \Delta y$ **2.89** **(0.02)**	$\Delta c \rightarrow \Delta cpi$ 1.90 (0.10)	$\Delta c \rightarrow ir$ **3.98** **(0.00)**	$\Delta c \rightarrow \Delta hp$ **2.58** **(0.04)**	$\Delta c \rightarrow \Delta m$ **6.31** **(0.02)**

Note: The table reports heteroskedasticity-robust test statistics for Ganger causality (F-tests). P-values are in parentheses. Significant test statistics are in bold.

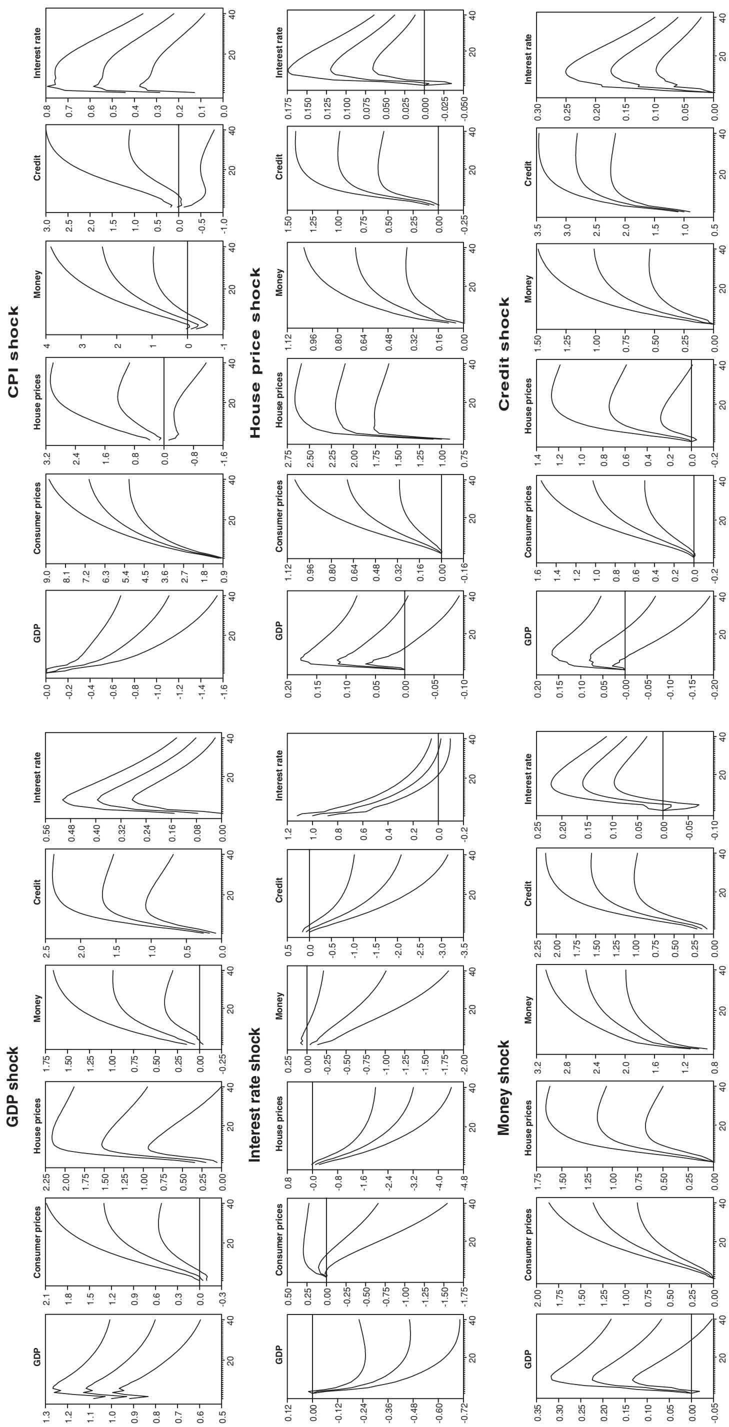
Standard errors for the impulse–response functions were computed based on a wild bootstrap (Goncalves and Kilian, 2004) in order to take into account potential heteroskedasticity of the residuals. The wild bootstrap is set up in the following way. For each draw, we first construct an artificial vector of innovations by multiplying each element of the vector of sample residuals $\varepsilon_{i,t}$ with an i.i.d. innovation drawn from the standard normal distribution. With these artificial innovations we construct artificial datasets based on the estimated VAR. The artificial dataset is then used to re-estimate the VAR and to generate shock-impulse responses based on the Choleski decomposition described above. This procedure is replicated 1,000 times and the bootstrapped standard error for each impulse–response function is obtained by calculating the standard error of the respective 1,000 shock-impulse responses.

IV. Empirical results

The panel VAR described in the previous section was first estimated over the longest possible sample period, 1973Q1–2006Q4, with a lag order of four, which was selected based on the Akaike information criterion. Table 1 displays the results from the Granger causality tests, which reveal strong evidence of multidirectional causality between house prices, money, credit, GDP, the CPI, and interest rates. In particular, monetary variables are found to have a significant effect on future house prices and, at the same time, house prices are found to have a highly significant effect on future money and credit growth. Monetary variables and house prices also significantly affect future GDP growth, while only money growth affects future CPI inflation.

Figure 3 displays the impulse responses to orthogonalized one-unit shocks in a two-standard-error band. Although, as we pointed out in the previous section, it is not possible

Figure 3: Impulse responses to orthogonalized one-unit shocks, sample 1973Q1–2006Q4



Note: The figures display impulse responses to a unit shock in a two-standard-error confidence band. Standard errors were calculated based on a wild bootstrap with 1,000 draws.

to attach a clear structural meaning to the orthogonalized shocks, the patterns of the impulse responses do in some cases allow a cautious structural interpretation. The dynamic effects of a GDP shock suggest that this shock mainly captures aggregate demand shocks: real GDP, consumer prices, and the nominal interest-rate increase. Also house prices and monetary variables increase both in nominal and real terms.[22] A CPI shock seems mainly to capture supply-side disturbances: the CPI increases, real GDP falls, the nominal interest rate increases and house prices, money, and credit fall in real terms. Also, the responses to an interest-rate shock are in line with prior expectations. Nominal interest rates increase temporarily, while all other variables fall. A house-price shock triggers significant increases in all the variables in the system. The same holds true for the dynamic effects of a money and of a credit shock. These results imply that there is a strong and highly significant multidirectional relationship between monetary variables, house prices, and the macroeconomy.

As we pointed out in the introduction, there are good reasons to conjecture that the dynamic link between house prices and monetary variables, but also the dynamics of the macroeconomy in general, changed in the early/mid-1980s owing to structural changes in financial systems and changes in the monetary policy regime. We would expect that the link between monetary variables and house prices has become stronger over the more recent sub-sample because of financial deregulation, while the reaction of consumer prices to macroeconomic shocks in general, and also shocks to money and credit, would be expected to have become weaker because of a more stability-orientated monetary policy. In order to test this hypothesis we replicate the empirical exercises for the sample period 1985Q1–2006Q4 and compare the results with those obtained from the full sample period.

Over this shorter sample period, the panel VAR was estimated with a lag order of three in accordance with the indications of the Akaike information criterion. Table 2 presents the results for the Granger causality tests. The test statistics reveal that the effects of monetary variables seem to have become weaker, while the effects of house prices seem to have become stronger. Many of the Granger causality tests for the monetary variables are now insignificant. Money and credit growth no longer have a significant effect on future output growth or on future house-price inflation. Money growth is also found not to Granger cause CPI inflation over the shorter sample period. Interestingly, credit growth now Granger causes CPI inflation, which was not the case over the full sample period. The Granger causality tests of the effects of house prices are all significant at least at the 5 per cent level, which is a stronger outcome than for the full sample period.

Figure 4 displays the impulse responses for the shorter sub-sample in a two-standard-error band together with the impulse–response function from the full sample (dotted lines) for comparison. Overall, the results are not fundamentally different from those obtained for the full sample. There is still clear evidence of a strong and highly significant multidirectional relationship between monetary variables, house prices, and the macroeconomy. There are, however, a number of interesting changes which are all in line with our prior expectations. First, the response of the CPI to the five different shocks is much weaker. In particular, while the CPI fell after some time following an interest-rate shock over the full sample, it now significantly increases. This may be interpreted as reflecting a more forward-looking

[22] The response of real house prices and real money and credit is given by the difference between the response of the respective nominal values and the response of the CPI. For example, if the CPI increases by more after a shock than nominal house prices, then real house prices fall.

Table 2: Granger causality tests (1985–2006 sample)

$\Delta y \rightarrow \Delta cpi$	$\Delta y \rightarrow ir$	$\Delta y \rightarrow \Delta hp$	$\Delta y \rightarrow \Delta m$	$\Delta y \rightarrow \Delta c$
1.19	**3.10**	**2.85**	1.10	**7.21**
(0.39)	**(0.03)**	**(0.04)**	(0.34)	**(0.00)**
$\Delta cpi \rightarrow \Delta y$	$\Delta cpi \rightarrow ir$	$\Delta cpi \rightarrow \Delta hp$	$\Delta cpi \rightarrow \Delta m$	$\Delta cpi \rightarrow \Delta c$
2.75	1.90	0.84	**2.77**	0.45
(0.04)	(0.13)	(0.47)	**(0.04)**	(0.71)
$ir \rightarrow \Delta y$	$ir \rightarrow \Delta cpi$	$ir \rightarrow \Delta hp$	$ir \rightarrow \Delta m$	$ir \rightarrow \Delta c$
1.89	**19.84**	**7.35**	1.44	0.76
(0.13)	**(0.00)**	**(0.00)**	(0.23)	(0.51)
$\Delta hp \rightarrow \Delta y$	$\Delta hp \rightarrow \Delta cpi$	$\Delta hp \rightarrow ir$	$\Delta hp \rightarrow \Delta m$	$\Delta hp \rightarrow \Delta c$
9.64	**3.35**	**3.75**	**8.83**	**11.58**
(0.00)	**(0.02)**	**(0.02)**	**(0.00)**	**(0.00)**
$\Delta m \rightarrow \Delta y$	$\Delta m \rightarrow \Delta cpi$	$\Delta m \rightarrow ir$	$\Delta m \rightarrow \Delta hp$	$\Delta m \rightarrow \Delta c$
1.27	0.84	**2.39**	1.51	0.71
(0.28)	(0.47)	**(0.07)**	(0.21)	(0.54)
$\Delta c \rightarrow \Delta y$	$\Delta c \rightarrow \Delta cpi$	$\Delta c \rightarrow ir$	$\Delta c \rightarrow \Delta hp$	$\Delta c \rightarrow \Delta m$
0.53	**7.52**	**4.50**	1.08	**5.98**
(0.66)	**(0.00)**	**(0.00)**	(0.36)	**(0.00)**

Note: The table reports heteroskedasticity-robust test statistics for Granger causality (F-tests). P-values are in parentheses. Significant test statistics are in bold.

conduct of monetary policy since the mid-1980s.[23] Second, the dynamic effects of a house-price shock on real GDP, the interest rate, and monetary variables have become stronger.[24] Third, the dynamic effects of a shock both to money and to credit have become weaker. However, since the response of the CPI has weakened by more than the response of nominal house prices, the effect of a shock to money or credit on *real* house prices has in fact become stronger. However, owing to the large confidence bands the difference between the full- and the sub-sample impulse responses is generally not statistically significant when the uncertainty surrounding both responses is taken into account.

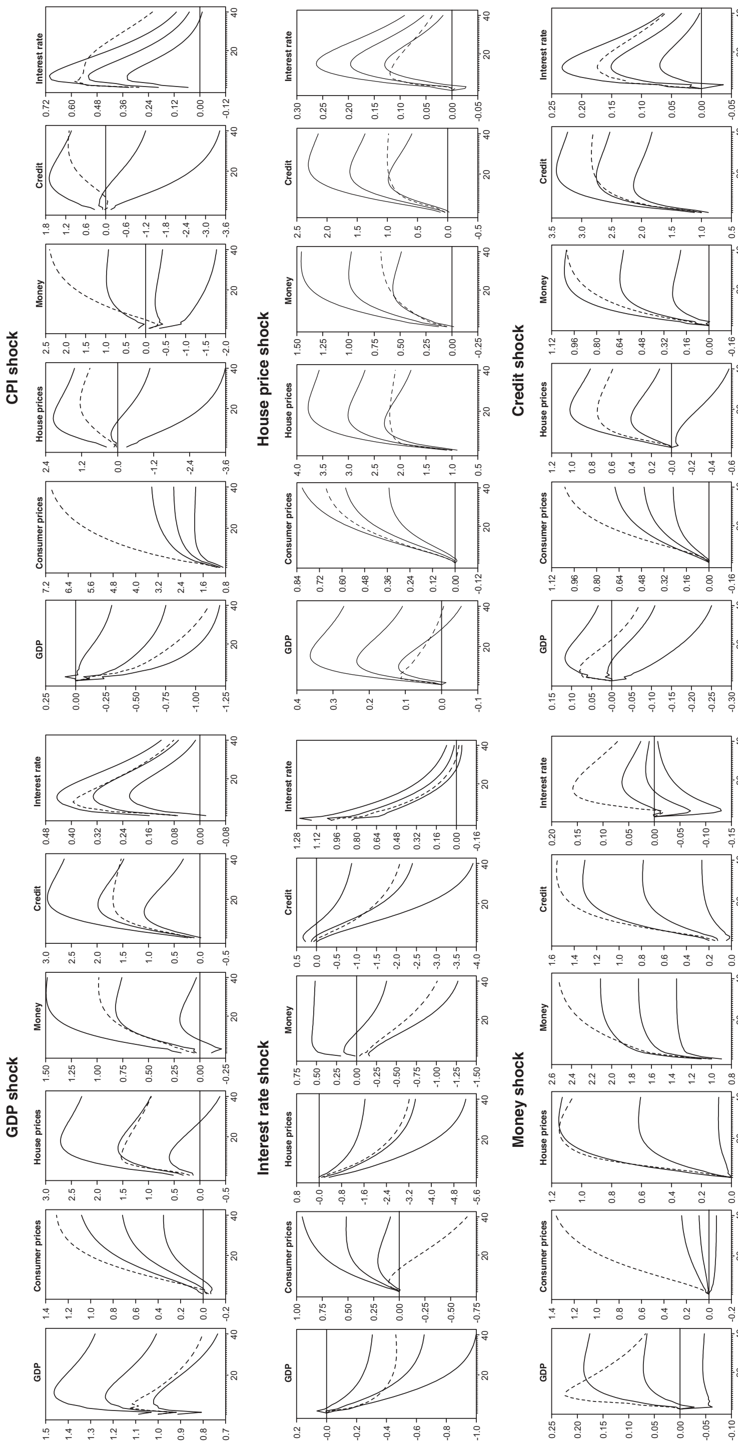
V. House-price booms

As we stated in the introduction, there is evidence that the link between monetary variables and asset prices is particularly close in times of asset-price booms. The panel set-up of the analysis also allows us to test the hypothesis that money and credit are more closely linked to

[23] The phenomenon of a positive response of the CPI to an interest-rate increase is known as the price puzzle and is attributed to forward-looking monetary policy, so that the impulse response captures in part also the reaction of monetary policy to expected future inflation.

[24] These results are consistent with evidence presented by Ludwig and Sløk (2004) and Muellbauer (2007), suggesting that the effect of house prices on consumption has become stronger since the mid-1980s.

Figure 4: Impulse responses to orthogonalized one-unit shocks, sample 1985Q1–2006Q4



Note: The figures display impulse responses to a unit shock in a two-standard-error confidence band. Standard errors were calculated based on a wild bootstrap with 1,000 draws. The dotted line is the impulse response from the full sample analysis.

Table 3: Episodes of house-price booms (1985–2006)

Australia	2001Q1–2005Q2
Belgium	1988Q3–2001Q1
Canada	1986Q3–1990Q1, 2002Q1–2006Q4
Denmark	1995Q1–2002Q2
Finland	1985Q1–1990Q3, 1997Q2–2006Q4
France	1988Q1–1991Q2, 2000Q1–2006Q4
Germany	1992Q1–1994Q4
Ireland	1995Q4–2004Q2
Italy	1989Q2–1993Q4, 2002Q2–2006Q4
Japan	1986Q4–1991Q1
Netherlands	1988Q2–2002Q3
Norway	1985Q1–1988Q3, 1995Q4–2002Q4, 2004Q1–2006Q4
Spain	1986Q4–1992Q2, 2002Q1–2006Q4
Sweden	1987Q4–1991Q4, 1998Q1–2006Q4
Switzerland	1986Q1–1990Q2, 2002Q1–2006Q4
UK	1986Q1–1990Q3, 1998Q3–2006Q2
USA	2000Q2–2006Q4

Note: The table reports periods of house-price booms. A boom is defined as a positive deviation of house prices from a smooth HP trend (smoothing parameter = 100,000) of more than 5 per cent lasting at least 12 quarters.

house prices when house prices are booming by running a dummy variable augmented panel VAR of the form:[25]

$$Y_{i,t} = A_i + A_{NB}(L)Y_{i,t} \times D_{i,t}^{NB} + A_B(L)Y_{i,t} \times D_{i,t}^{B} + \varepsilon_{i,t} \tag{3}$$

where $D_{i,t}^{B}$ is dummy variable that is set equal to one when there is a house-price boom in period t in country i, and equal to zero otherwise. $D_{i,t}^{NB}$ is, in turn, a dummy variable that is set equal to one when there is no house-price boom in period t in country i, and equal to zero otherwise.

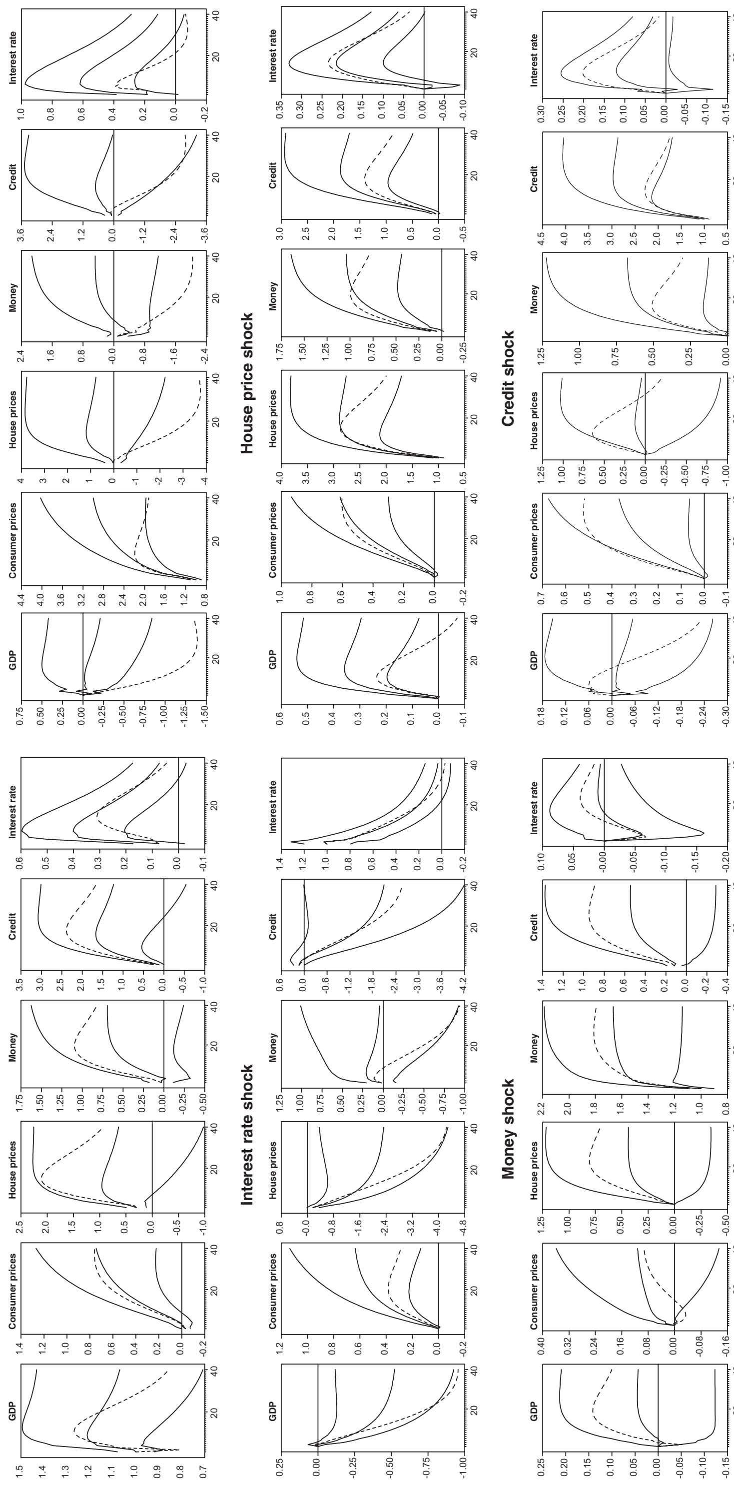
In light of the results in the previous section, our analysis focuses on the more recent sample period, 1985–2006. Following the approach of Borio and Lowe (2004) and Adalid and Detken (2007) to define aggregate asset-price booms, we define a house-price boom as a persistent deviation of real house prices from a smooth trend, calculated based on a one-sided Hodrick–Prescott (HP) filter with a smoothing parameter of 100,000. A boom is defined as a positive deviation of house prices from this smooth trend of more than 5 per cent lasting at least 12 quarters. The boom episodes identified in this way are reported in Table 3.

The dummy augmented panel VAR in (3) was estimated with the dummy variables specified in line with the identified boom episodes. Figure 5 displays the impulse responses. The solid lines are the impulse responses (in a two-standard-error band) for the no-boom scenario, i.e. obtained from a simulation of the estimated panel VAR with the no-boom dummy $D_{i,t}^{NB}$, set equal to one, and the boom dummy set equal to zero. The dotted lines are the impulse responses obtained under the boom scenario, i.e. when the impulse responses are simulated with the dummy $D_{i,t}^{B}$ set equal to one and the dummy $D_{i,t}^{NB}$ set equal to zero.

The results reveal that the dynamic repercussions of most shocks are stronger during house-price booms. In particular, the effect of money and credit shocks on the economy and on

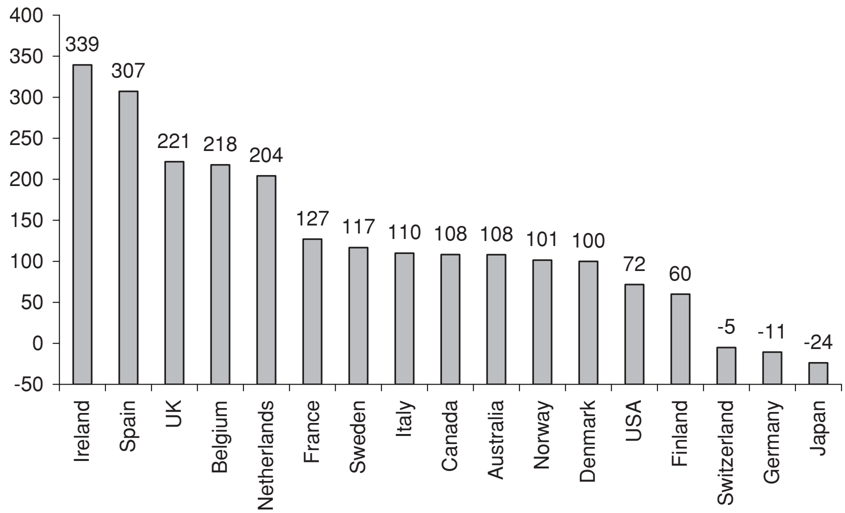
[25] In a recent paper, Adalid and Detken (2007) have shown, also using a panel framework, that broad money growth has a particularly strong influence on real house-price growth when there is an aggregate asset-price boom, where the aggregate asset price is a BIS construct aggregating share prices, residential property prices, and commercial property prices weighted by their respective share in household wealth.

Figure 5: Dynamics during house-price booms (1985–2006)



Note: The dotted line is the impulse response for episodes of house-price booms. The solid lines are the responses in a two-standard-error band when there is no boom.

Figure 6: Accumulated real house-price increases 1985–2006 (in %)



nominal and real house prices is stronger.[26] This result supports the view that money and credit growth contain useful information about emerging house-price booms or bubbles. Somewhat surprisingly, the dynamic effects of house-price shocks on the other variables in the system are not stronger during house-price booms. However, if the uncertainty surrounding both impulse responses is taken into account, the difference between any two boom- and no-boom-impulse responses is in general not statistically significant owing to the large confidence bands.

An alternative way to test the boom hypothesis is to assess whether the link between monetary variables and house prices has been stronger in countries that have experienced particularly strong house-price increases over the sample period. Figure 6 shows the accumulated increases in real house prices (in%) over the period 1985–2006 in each of the 17 countries covered by this study. There is a group of five countries (Ireland, Spain, the UK, Belgium, and the Netherlands) which has experienced very strong real house-price increases of more than 200 per cent over this period. The majority of countries experienced a moderate increase in real house prices of between 60 and 130 per cent. There are three countries (Switzerland, Germany, and Japan) where real house prices decreased over this sample period.

In order to test whether the dynamic interaction between house prices and monetary variables is different in countries which have experienced relatively strong real house-price increases, we re-estimate the dummy variable augmented panel VAR in (3), but now specified in order to separate the countries with particularly high real house-price inflation since 1985

[26] These results are broadly consistent with those reported by Adalid and Detken (2007). However, while they find that only money growth influences future house prices when there is an asset-price boom, we find that both money and credit matter. There are three potential explanations for this discrepancy. First, Adalid and Detken (2007) focus on the effect of monetary variables on real house prices in a single-equation framework, while our analysis is based on a multivariate framework. Second, Adalid and Detken investigate the link between house prices and money and credit during aggregate asset-price booms (see footnote 25), while we investigate the dynamics during house-price booms. Finally, there is a difference in sample periods. The sample period in Adalid and Detken is 1972–2004, while here it is 1985–2006.

Table 4: Typical or maximum LTVs in industrialized countries (in %)

	1980s	1990s	Recent
Australia	80	80	60–70
Belgium	75	80	80–85
Canada	75	80	75–95
Denmark	95	80	80
Finland	85	70–80	70–85
France	80	70–80	66
Germany	65	60–80	70
Ireland	80	80	91–95
Italy	56	40	80
Japan	60	–	70–80
Netherlands	75	75	112
Norway	80	80	60–80
Spain	80	70–80	83
Sweden	95	70–75	90
Switzerland	–	–	65–80
UK	87	90–95	80
USA	89	89	85

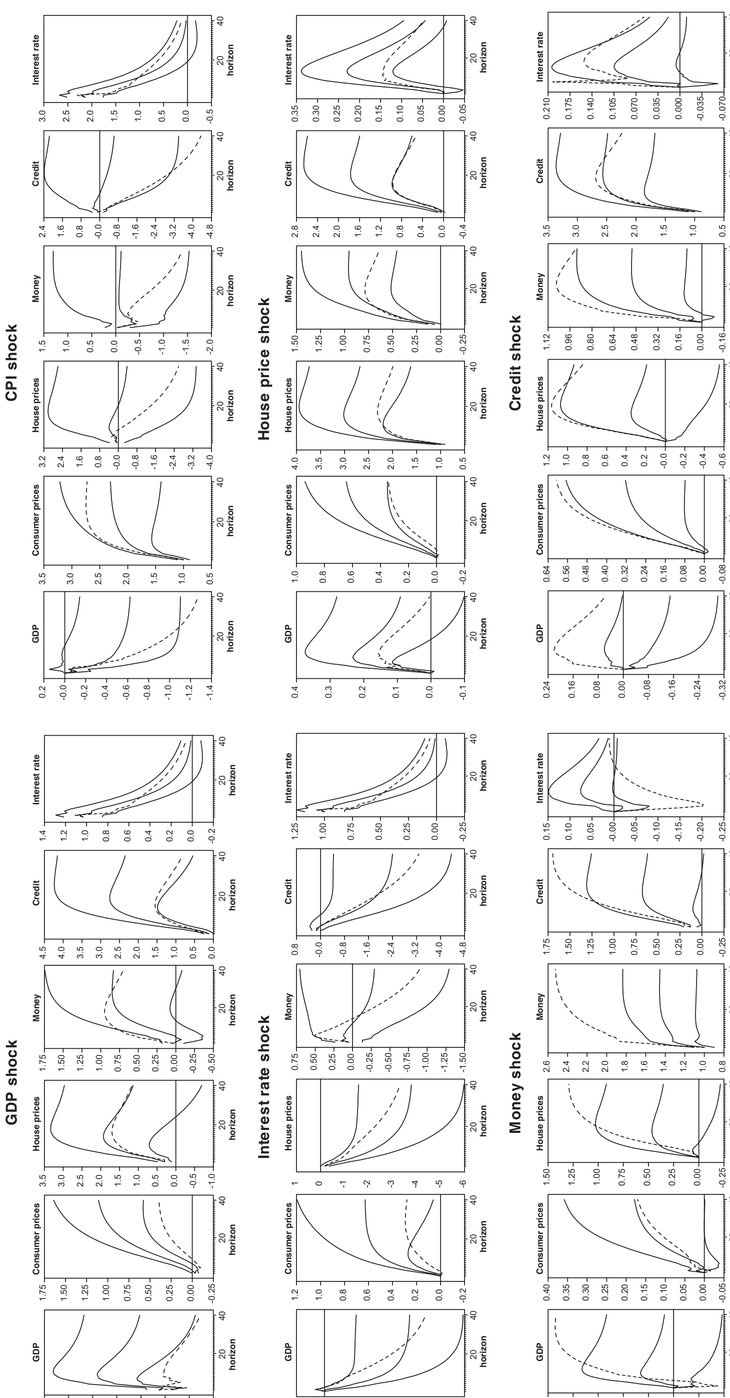
Sources: Japelli and Pagano (1994), Maclennan *et al*. (2000), Chiuri and Japelli (2003), OECD (2004), BIS (2006), Miles and Pillonca (2007).

from the other countries. This was achieved by setting the dummy variables $D_{i,t}^{B}$ equal to one for the five countries with high real house-price increases and equal to zero for all other countries, while the dummy $D_{i,t}^{NB}$ was specified conversely.

Figure 7 displays the impulse responses. The solid lines are the impulse responses (in a two-standard-error band) for the countries with moderate and low house-price increases, i.e. obtained from a simulation of the estimated panel VAR with the dummy $D_{i,t}^{NB}$ set equal to one and the boom dummy set equal to zero. The dotted lines are the impulse responses obtained for the high house-price-increase countries, i.e. when the impulse responses are simulated with the dummy $D_{i,t}^{B}$ set equal to one and the dummy $D_{i,t}^{NB}$ set equal to zero. The results again support the view that money and credit growth are useful indicators of house-price booms. Money and credit shocks are found to have a stronger effect on nominal and real house prices in the group of countries characterized by high real house-price inflation. The dynamic effects of house-price shocks on the other variables in the system are not found to be stronger in these countries, confirming the finding of the previous exercise that house-price movements do not have stronger repercussions on the economy during periods of a house-price boom.

As a final exercise, we take a brief look at cross-country differences in household borrowing constraints and how they evolved over time. Household borrowing against real-estate collateral is usually restricted by a wealth constraint, a loan-to-value ratio (LTV) restricting the extended loan from exceeding a certain proportion of the value of the house, and/or an income constraint restricting mortgage interest payments from exceeding a certain proportion of the borrower's income. While cross-country data on the latter are generally not available, data on maximum or typical LTVs are available, though only pointwise. Based on various sources, we collected three observations on typical or maximum LTVs broadly referring to the 1980s, the 1990s, and the more recent period, which are reported in Table 4.

Figure 7: High house-price-inflation countries versus moderate/low house-price-inflation countries (1985–2006)



Note: The dotted line is the impulse response for the countries which experienced particularly high rates of house price inflation over the sample period (Belgium, the Netherlands, Ireland, Spain, and the UK). The solid lines are the responses in a two-standard-error band for the remaining countries.

These figures reveal that the five countries with the highest house-price increases since 1985 were also characterized by relatively high LTVs, at least more recently. The figures also show, however, that there is no perfect correlation between house-price increases and LTVs. For example, LTVs in the USA have historically also been very high, but the 70 per cent house-price increase since 1985 was relatively moderate. There also seems to be a link between LTVs and the house-price boom episodes reported in Table 3. For example, LTVs in Sweden, Norway, and Finland were high in the 1980s when these countries experienced house-price booms, which were followed by banking crises. However, there have also been many house-price booms occurring in countries and at times when LTVs were moderate or low.

VI. Conclusions

The empirical analysis of this paper offers a number of interesting insights. There is evidence of a significant multidirectional link between house prices, broad money, private credit, and the macroeconomy. Money growth has a significant effect on house prices and credit, credit influences money and house prices, and house prices influence both credit and money. This link is found to be stronger over a more recent sub-sample from 1985 to 2006 than over a longer sample going back to the early 1970s, a finding that most likely reflects the effects of financial system liberalizations in industrialized countries during the 1970s and early 1980s. Furthermore, shocks to house prices, credit, and money all have significant repercussions on economic activity and aggregate price inflation. Shocks to GDP, the CPI, and the interest rate are in turn found to have significant effects on house prices, money, and credit. The empirical analysis further reveals that the effects of shocks to money and credit on house prices are stronger when house prices are booming than otherwise.

These findings suggest that a monetary policy strategy that gives due weight to the analysis of monetary developments could, in principle, induce the central bank to react indirectly to emerging asset-price bubbles and thereby mitigate adverse longer-run consequences of financial imbalances. However, in times of low and stable inflation, central banks might find it difficult to communicate such a leaning-against-the-wind policy. In a currency union like the euro area, there is the further problem of regional differences in house-price and credit dynamics, which can only be addressed by a common monetary policy to the extent that they are reflected in the area-wide aggregates (Goodhart, 2005). A way out of these dilemmas might be to consider a secondary financial instrument that could directly address the link between house prices and monetary variables, and could also be used at the regional level in a currency union. In previous work (Goodhart and Hofmann, 2004*b*, 2007) we have made the suggestion that LTVs on mortgage lending should be varied countercyclically. Thus the LTV could be raised when mortgage growth (and house-price inflation) was low or declining, and lowered during booms. Measures of this kind have been applied in the past in Hong Kong and South Korea and, more recently, in Estonia.

The case for such a secondary instrument is, however, not empirically supported by the analysis of this paper. Based on pointwise descriptive analysis, we have uncovered only a weak correlation between the level of LTVs and cross-country differences in house-price increases or episodes of house-price booms. But this might well be due to the fact that the quality of cross-country data on maximum or typical LTV levels is rather poor and that other structural features of mortgage markets also play a role for which it is difficult to control. A more rigorous empirical and theoretical analysis of the role of the level of LTVs for house-price and monetary dynamics and their interactions would, in our view, be a fruitful avenue for future research.

Appendix

Table A1: Details on the house-price data series

Australia	Price index of established houses, weighted average of eight state capital cities (Sydney, Melbourne, Brisbane, Adelaide, Perth, Hobarth, Darwin, Canberra); quarterly data 1970Q1–2006Q4 (source: BIS)
Belgium	Price index of existing and new dwellings; quarterly data 1980Q1–2006Q4 (source: BIS); annual data 1970–1979 (source: BIS) interpolated based on the Chow–Lin procedure using a construction cost index (source: BIS) as reference series
Canada	Average prices of existing homes; quarterly data 1970Q1–2006Q4 (source: BIS)
Denmark	Price index of new and existing houses, good and poor condition; quarterly data 1971Q1–2006Q4 (source: ECB)
Finland	Price index of new and existing dwellings; quarterly data 1978Q1–2006Q4 (source: BIS); annual data 1970–1977 (source: BIS) interpolated based on the Chow–Lin procedure using the rent CPI (source: OECD MEI) as reference series
France	Price index for existing dwellings; quarterly data 1996Q1–2006Q4 (source: ECB); price index for existing homes; annual data 1970–1995 (source: BIS) interpolated based on the Chow–Lin procedure using for 1980Q2–1995Q4 a price index for existing flats in Paris (source: ECB) and for 1970Q1– 1980Q1 a cost index for new residential construction (source: BIS) and the rent CPI (source: OECD MEI) as reference series
Germany	Prices of good-quality existing dwellings in 125 cities (in four large cities prior to 1975); annual data 1970–2006 (source: BIS) interpolated based on the Chow–Lin procedure using a building cost index (source: BIS) and the rent CPI (source: OECD MEI) as reference series
Ireland	Second-hand house prices (from 1978) and new house prices (prior to 1978); quarterly data 1975Q1–2006Q4 (source: Irish Department of the Environment); new house prices; annual data 1970–1974 (source: ECB) interpolated based on the Chow–Lin procedure using the rent CPI (source: OECD MEI) as reference series
Italy	Price index of new and existing dwellings; semi-annual data (source: ECB) interpolated based on the Chow–Lin procedure using a construction cost index (source: BIS) and the rent CPI (source: OECD MEI) as reference series
Japan	Residential land price index; semi-annual data (source: BIS) interpolated based on the Chow–Lin procedure using the housing investment deflator (source: OECD QNA) and the rent CPI (source: OECD MEI) as reference series
Netherlands	Price index for one-family houses and existing flats; quarterly data 1970Q1–2006Q4 (source: BIS)
Norway	Registered purchase price of all dwellings; quarterly data 1970Q1–2006Q4 (source: BIS)
Spain	Price index of new and existing dwellings; quarterly data 1987Q1–2006Q4 (source: BIS); Madrid house prices; annual data 1971–86 (source: BIS) interpolated based on the Chow–Lin procedure using a construction cost index (source: OECD MEI) and the rent CPI (source: OECD MEI) as reference series
Sweden	Price index of owner-occupied new and existing dwellings; quarterly data 1970Q1–2006Q4 (source: BIS)
Switzerland	Price index of single-family homes and owner-occupied flats; quarterly data 1970Q1–2006Q4 (source: BIS)
UK	Price index of new and existing dwellings; quarterly data 1970Q1–2006Q4 (source: BIS)
USA	Price index of existing homes; quarterly data 1970Q1–2006Q4 (source: BIS). The BIS series links the Office of Federal Housing Enterprise Oversight (OFHEO) house-prices index to the National Association of Realtors' house-price index in 1975Q1

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