Bank Lending and Property Prices: Some International Evidence*

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(PRELIMINARY AND INCOMPLETE)

ABSTRACT

Over the last two decades, bank lending has been closely correlated with property prices in both industrialised and developing countries. From previous empirical studies it is not clear whether these cycles were driven by property prices or lending. In this study I analyse the direction of causality between bank lending and property prices in 16 industrialised countries over the last two decades. Long-run causality appears to go from property prices to bank lending. This finding suggests that property price cycles, reflecting changing beliefs about future economic prospects, drive credit cycles, rather than excessive bank lending being the cause of property price bubbles. The results of estimating error-correction equations for credit growth and the change in property prices suggest that there is evidence of short-run causality going in both directions. This implies that a mutually re-enforcing element in past boom bust cycles in credit and property markets cannot be ruled out.

*This paper was written while the author was a Visiting Research Fellow at the Hong Kong Institute for Monetary Research at the Hong Kong Monetay Authority. I am grateful to Claudio Borio, Stefan Gerlach, Olan Henry, Charles Goodhart, Philip Lowe, Jürgen von Hagen and seminar participants at the Bank for International Settlements for helpful comments and Steve Arthur for his help with the data. All remaining errors are my own responsibility.

1. Introduction

Over the last decades, bank lending was closely correlated with property prices in both industrialised and developing countries. The coincidence of cycles in credit and property markets has been widely documented in the policy oriented literature (IMF, 2000, BIS, 2001), but there are only few studies trying to disentangle the direction of causality between bank lending and property prices.

Property prices may affect bank lending via various wealth effects. Households and firms may be borrowing constrained due to asymmetric information in the credit market, which gives rise to adverse selection and moral hazard problems. As a result, households and firms can only borrow when they can offer collateral, so that their borrowing capacity is a function of their collateralisable net worth[1]. Since property is commonly used as collateral, property prices are an important determinant of the private sector's borrowing capacity.

Data on the composition of household wealth reported in OECD (2000) reveal that households hold a large share of their wealth in property. A change in property prices may therefore have a large effect on consumers' perceived lifetime wealth, inducing them to change their spending and borrowing plans ad thus their credit demand in order to smooth consumption over the life cycle[2].

Finally, property prices affect the value of bank capital, both directly to the extent that banks own assets, and indirectly by affecting the value of loans secured by assets[3]. Via their effect on banks' balance sheets, property prices influence the risk taking capacity of banks and thus their willingness to extend loans.

Bank lending may affect property prices via various liquidity effects. The price of property may be seen as an asset price, which is determined by the discounted future stream of

[1] Basic references of this literature are Bernanke and Gertler (1989) and Kiyotaki and Moore (1997). For a survey see Bernanke, Gertler and Gilchrist (1998). An early reference is Fisher (1933).
[2] The lifecycle model of household consumption was originally developed by Ando and Modigliani (1963). A formal exposition of the lifecycle model can be found in Deaton (1992) and Muellbauer (1994).

property returns. An increase in the availability of credit may lower interest rates and stimulate current and future expected economic activity. As a result, property prices will rise because of higher expected returns on property and a lower discount factor.

Property may also be seen as a durable good in temporarily fixed supply. An increase in the availability of credit may increase the demand for housing if households are borrowing constrained. With supply temporarily fixed because it takes time to construct new housing units, this increase in demand will be reflected in higher property prices.

Theory therefore suggests that causality between bank lending and property prices may go in both directions. This two-way causality may give rise to mutually reinforcing cycles in credit and property markets[4]. A rise in property prices, caused by more optimistic expectations about future economic prospects, raises the borrowing capacity of firms and households by increasing the value of collateral. Part of the additional available credit may also be used to purchase property, pushing up property prices even further, so that a self-reinforcing process can evolve.

Little empirical research has been done on the relationship between credit and asset prices. Most studies rely on a single equation set up, focusing either on bank lending or property prices. Goodhart (1995) finds that property prices significantly affect credit growth in the UK but not in the US. Hilbers, Lei and Zacho (2001) find that the change in residential property prices significantly enter multivariate probit-logit models to explain the outbreak of financial distress in industrialised and developing countries. Borio and Lowe (2002) show that a measure of the aggregate asset price[5] gap, measured as the deviation of aggregate asset prices from their long-run trend, combined with a similarly defined credit gap measure, is a useful indicator of financial distress in industrialised countries.

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[3] Chen (2001) develops an extension of the Kiyotaki and Moore (1997) model where an additional amplification of business cycles results from the effect of asset price movements on banks‘ balance sheets. An early reference for this argument is Keynes (1931).

[4] The possibility of mutually reenforing cycles in credit and asset markets in general has already been stressed by Kindleberger (1978) and Minsky (1982).

[5] Aggregate asset price indices are calculated as a weighted average of residential property prices, commercial property prices and equity prices. The weights are based on the share of each asset in national balance-sheets, which are derived based on national flow-of-funds data or UN standardised national accounts. The index weight of both residential and commercial property prices is on average above 80% so that property price movements dominate the movements of the aggregate asset price index.

Borio, Kennedy and Prowse (1994) investigate the relationship between credit to GDP ratios and aggregate asset prices[6] for a large sample of industrialised countries over the period 1970-1992 using annual data. They focus on the determinants of aggregate asset price fluctuations, hypothesising that the development of credit conditions as measured by the credit to GDP ratio can help to explain the evolution of aggregate asset prices. 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 - early 1990s would have been much less pronounced or would not have occurred at all had credit ratios remained constant. For a panel of four East Asian countries (Hong Kong, Korea, Singapore and Thailand), Collyns and Senhadji (2001) find that credit growth has a significant contemporaneous effect on residential property prices. They conclude that bank lending has contributed significantly to the real estate bubble in Asia prior to the 1997 East Asian crisis.

Al these studies potentially suffer from simultaneity problems and cannot disentangle the direction of causality between credit and property prices. In two recent studies, Gerlach and Peng (2002) and Hofmann (2001) analyse the relationship between bank lending and property prices based on a multivariate empirical framework. Gerlach and Peng (2002) find that in Hong Kong, both long-run and short-run causality goes from property prices to lending, rather than conversely. Hofmann (2001) finds for a set of 16 industrialised that including property prices in the empirical model is decisive to explain the long-run development of bank lending and that long-run causality goes from property prices to bank lending. Based on the analysis in Hofmann (2001), I will in the following assess the pattern of long-run and short-run causality between bank lending and property prices for a set of 16 countries. Based on Johansen's (1988) approach to cointegration analysis I first test for cointegration between bank lending, real GDP and property prices. Based on error-correction models I then test for the pattern of long-run and short-run causality between bank lending and property prices.

The plan of the paper is as follows. The following Section 2 describes the data used for the empirical analysis and presents some international stylised facts about the comovement of

[6] They use the same measure of aggregate asste prices as Borior and Lowe (2002).

bank lending and property prices over the period 1984-2001. Section 3 describes the empirical methodology used to test for the presence of long-run relationships between ecnomic activity, property prices and bank lending and presents the test results and the estimated long-run relationships. In Section 5 I estimate error-correction models and test for long-run and short-run causality between bank lending and property prices. Section 5 concludes.

2. Bank Lending and Property Prices: Data and Stylised Facts

In the following sections I analyse the relationship between aggregate bank lending, aggregate economic and residential property prices in 16 countries: the US, Japan, Germany, France, Italy, the UK, Canada, Australia, Sweden, Norway, Finland, the Netherlands, Belgium, Ireland, Hong Kong and Singapore.

The bank lending series used corresponds to the total credit aggregate to the non-bank private sector from the Banking Survey in the IMF International Financial Statistics. Nominal credit aggregates were transformed into real terms by deflation with the consumer price index. As a measure of aggregate economic activity I use real GDP. Quarterly residential property price indices were available for all countries except for Japan, Italy and Germany. For Japan and Italy semi-annual indices were transformed to quarterly frequency by linear interpolation. For Germany a quarterly series was generated by linear interpolation based on annual observations from the first quarter of each year. In order to obtain a measure of real property prices, nominal property prices were deflated with the consumer price index. In Section 3 I also allow a short-term real interest rate to enter the error-correction equations, using an ex-post measure of the short-term real interest rate, measured as the three months interbank money market rate[7] less quarterly CPI inflation. The data for the industrialised countries were

[7] A more accurate measure of aggregate financing costs would of course be an aggregate lending rate. Representative lending rates are, however, not available for most countries. Empirical evidence suggests that short-term and long-term lending rates are in the long-run tied to money market rates or policy rates (see Borio and Fritz (1995) for a large sample of industrialised countries, Hofmann (2001) for euro area countries and Hofmann and Mizen (2001) for the UK), so that money market rates appear to be useful approximations of the financing costs of credit.

taken from the BIS database. Data for Hong Kong and Singapore are from the CEIC database. Except for the nominal interest rate, all data were seasonally adjusted using the Census X-12 procedure.

Figure 1 displays the year-on-year growth rates of real GDP (thick solid line), bank lending (thin solid line) and residential property prices (dotted line). The graphs reveal that, over the last two decades, all countries in our sample have experienced at least one boom bust cycle in bank lending. Following the liberalisation and deregulation of credit markets in the early-mid 1980s[8], most countries in our sample experienced a boom bust cycle in bank lending in the late 1980s – early 1990s. The cycles were particularly violent in the Nordic countries, where, in the wake of the banking crisis, bank lending declined by around 30% in Sweden and Norway and by almost 50% in Finland compared to its previous peak[9].

The Japanese banking crisis in the 1990s is also reflected by a sharp drop in bank lending. What makes the Japanese crisis different from the experience of the other countries in our sample is the long duration of the crisis. While in all other countries, including the Nordic countries, bank lending recovered in the early-mid 1990s, Japan's banking sector problems continued throughout the 1990s up to the present day[10].

In the late 1990s, most countries experienced again a lending boom, which was particularly marked in Ireland and the Netherlands with double digit growth rates in bank lending. Hong Kong and Singapore experienced a lending boom in the early-mid 1990s followed by a marked bust in 1997/98 in the wake of the East Asian crisis.

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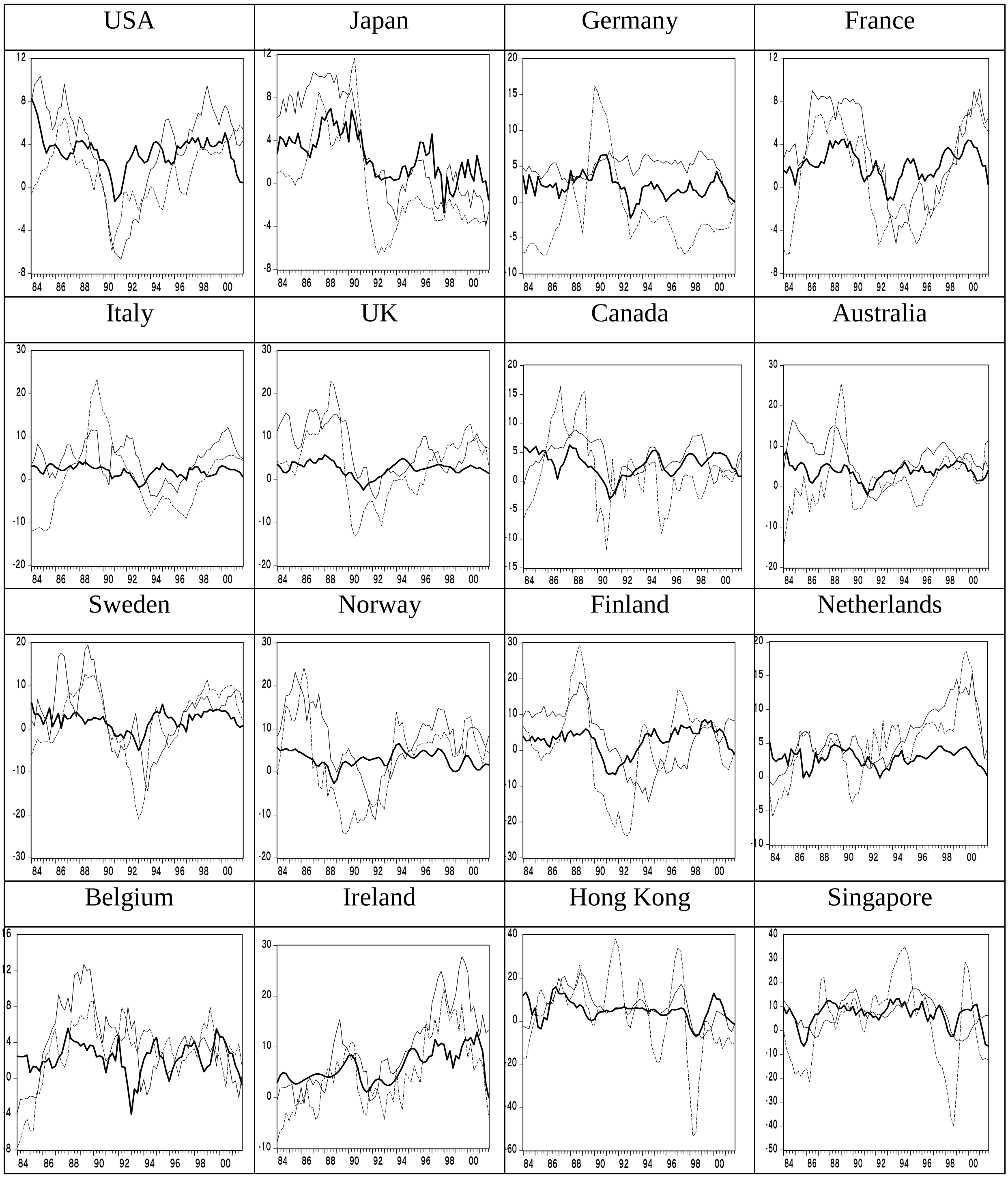
[8]See BIS (1999) for a compilation of articles reviewing the development of financial sectors in industrialised countries since the 1980s.

[9] Drees and Pazarbasioglu (1998) review the causes and policy implications of the Nordic banking crises.

[10] The literature on the Japanese banking crisis is of course enormous. See Hoshi and Kashyap (1999) for a recent survey and the references therein.

Figure 1: Bank lending, economic activity and property prices



Source: BIS, IMF, CEIC, national sources.. The dotted line represents credit growth (right hand scale) and the solid line represents the rate of change in the residential house price index (left hand scale).

The graphs reveal that bank lending is closely correlated with real GDP and property prices[11]. It appears that the development of bank lending coincides with or follows the development of real GDP, while property prices appear to lead bank lending and economic activity. This observation suggests that bank lending adjusts to real economic activity and expectations about future economic prospects reflected in property prices, rather than excessive bank lending, in the wake of financial liberalisation, being the source of business cycles and property price bubbles. In the following sections I will investigate whether this visual impression is supported by formal econometric analysis.

3. Long-Run Relationships

Standard augmented Dickey-Fuller (Dickey and Fuller, 1981) and Phillips-Perron (Phillips and Perron, 1988) unit root tests suggest that the natural logs of real bank lending, real property prices and real GDP are integrated of order one in all countries under investigation. The short-term real interest rate appears to be a borderline case. The null of non-stationarity can be rejected in about half of the countries. In the other countries the test statistic is on the margin of significance. Since there are strong theoretical reasons to believe that the real interest rate is stationary[12], I assume in the following analysis that the short-term real interest rate is I(0). In the following analysis of long-run relationships I therefore do not include the real interest rate in the empirical model but enter it in levels in the error-correction models estimated in Section 4.

The analysis of long-run relationships between real lending, real GDP and real property prices is based on the multivariate approach to cointegration analysis proposed by Johansen (1988).

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[11] The coincidcence of these cycles has previously been been extensively documented in the the policy oriented literature (Borio, Kennedy and Prowse, 1994, IMF, 2000, BIS, 2001, Borio and Lowe, 2002).

[12] The intetemporal Euler consumption equation implies that the long-run level of the real interest rate is equal to the sum of constant time preference rate and the constant long-run growth rate of consumption.

In order to test for cointegration I estimate for each country a VECM model of the form:

(1) $$\Delta x_{t} = C_1 \Delta x_{t-1} + ... + C_{k-1} \Delta x_{t-k} + C_0 x_{t-1} + \mu + \varepsilon_t ,$$

where *x* is a vector of endogenous variables comprising the log of real bank lending, real GDP and real property prices. μ is a vector of constants and ε is a vector of white noise error terms. Since I want to allow for deterministic time trends in the levels of the data I leave the constant unrestricted. The sample period for the analysis is the first quarter 1984 till the fourth quarter of 2001. Allowing for a maximum lag length of ten, the lag order *k* of the VECMs was determined by consulting the Akaike and the Schwarz-Bayes information criterion and Likelihood ratio tests of lag order reduction and then choosing the most parsimonious specification that is consistent with non-autocorrelated and non-heteroskedastic residuals.

The Johansen methodology is based on maximum likelihood estimation and aims at testing the rank of the matrix C_0, which indicates the number of long-run relationships between the endogenous variables in the system[13]. In Table 1 I report the test statistics of the Trace test for cointegration[14], together with the lag specification of the VECMs and a system Lagrange Multiplier test for autocorrelation up to order five, a system White test for heteroskedasticity and a system Jarque-Berra test for normality of the residuals.

The diagnostic tests suggest that there is no evidence of serial correlation or heteroskedasticity. In some cases there is evidence of non-normality of the residuals, which is in all cases due to excess kurtosis of the residuals and does therefore not invalidate the analysis (Johansen, 1995). Recursive Chow breakpoint tests (not reported) suggest that the estimated systems are stable over sub-samples.

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[13] For a detailed technical exposition of the Johansen approach see Johansen (1995).
[14] We do not report the results of Maximum Eigenvalue test since it is inferior to the Trace test (Johansen, 1994). The results of the Maximum Eigenvalue test were in general found to be consistent with the results of the reported Trace test.

In order to assess the significance of the Trace test statistics I use critical values tabulated in Osterwald-Lennum (1992). Taking the 5% significance level as the reference value, the Trace test for cointegration suggests for eleven out of the sixteen countries the presence of a single long-run relationship. For Sweden and Norway the test suggests two long-run relationships. For Japan, Canada, and Belgium the Trace test does reject the null of no cointegration at the five percent level. The test statistic is for these countries, however, still higher than or very close to the 10% critical value of 27.85.

The cross-country evidence therefore strongly suggests the presence of a single long-run relationship between bank lending, GDP and property prices[15]. Given that, the matrix C_0 can be factorised as $C_0 = \alpha\beta'$. α is a (3*x1*) vector of loading or adjustment coefficients and β is a (3*x1*) vector of cointegrating or long-run coefficients. The cointegrating coefficients β describe the relationship linking the endogenous variables in the long-run. The loading coefficients α describe the dynamic adjustment of the endogenous variables to long-run equilibrium given by $\beta' x$.

The cointegrating vectors were identified by normalising on bank lending. If supported by the data, further overidentifying homogeneity or zero restrictions were imposed on the β and α vectors. The estimated long-run and loading coefficients and a likelihood ratio test of the imposed overidentifying restrictions are shown in Table 2. Asymptotic standard errors for the long-run and the loading coefficients and probability values for the likelihood ratio test are reported in parentheses.

[15] We leave the second cointegrating relationship indicated for Sweden and Norway unidentified in order to achieve a consistent modelling of the long-run and short-run dynamics across countries. Closer investigation

Table 1: Cointegration Test Results

	Cointegration Test			*Diagnostics*		
	$r = 0$	$r \leq 1$	$r \leq 2$	*SC*	*H*	*N*
USA Lags = 6	31.72*	9.97	1.78	9.82	234.97	19.72**
Japan Lags = 3	27.89	13.6	1.39	4.34	100.96	7.66
Germany Lags = 4	39.29**	13.96	1.57	7.02	142.10	149.50**
France Lags = 1	30.12*	7.61	0.43	6.98	57.44	4.48
Italy Lags = 10	39.94**	6.70	0.92	7.21	379.12	72.13**
UK Lags = 1	30.63*	10.90	2.21	7.16	205.30	15.52*
Canada Lags = 3	27.65	10.30	0.16	8.10	134.96	13.98*
Australia Lags = 4	31.40*	10.13	2.01	3.26	154.61	12.21
Sweden Lags = 8	54.96**	18.17*	0.17	7.47	323.83	29.13**
Norway Lags = 2	66.84**	30.99**	4.29*	9.49	103.70	4.75
Finland Lags = 4	31.48*	12.48	0.01	11.28	162.62	8.07
Netherlands Lags = 8	30.32*	5.44	0.24	7.65	293.58	20.33**
Belgium Lags = 5	29.61	10.39	0.03	2.51	184.53	14.51*
Ireland Lags = 9	45.58**	12.26	4.20*	9.62	359.69	32.29**
Hong Kong Lags = 7	40.19**	7.56	0.11	8.96	283.02	17.04**
Singapore Lags = 4	31.29*	6.10	0.41	7.30	49.48	4.34

*Note: The table displays the test statistics of the Johansen trace test for cointegration. The 5% (1%) critical values for the cointegration test are 29.68 (35.65), 15.41 (20.04), 3.76 (6.65) for r=0, r=1, r=2 respectively (Osterwald-Lenum, 1992). SC is a system Lagrange multiplier test for residual serial correlation up to order 5, H is a system White test for residual heteroskedasticity and N s a system Jarque-Berra test for residual normality. * and ** indicates significance of a test statistic at the 5% and 1% level respectively.*

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suggested that the second cointegrating vector might be a long-run asset pricing relationship linking property prices to real GDP.

Table 2: Long-run Relationships

	Long-run coefficients			Loading coefficients			Restriction
	Bank lending	Real GDP	Property Prices	Bank lending	Real GDP	Property Prices	
USA	1.0	-0.623 (0.033)	-1.0	-0.112 (0.027)	0	0	1.57 (0.67)
Japan	1.0	-1.0	-0.604 (0.092)	-0.10 (0.03)	0	0	2.30 (0.51)
Germany	1.0	-2.285 (0.084)	0	-0.024 (0.015)	0.08 (0.021)	0	1.47 (0.48)
France	1.0	-0.79 (0.062)	-1.0	-0.16 (0.033)	0	0	0.41 (0.94)
Italy	1.0	-2.083 (0.08)	0	-0.183 (0.053)	0	0	4.89 (0.18)
UK	1.0	-1.606 (0.028)	-0.633 (0.135)	-0.087 (0.028)	0.03 (0.012)	0	1.14 (0.29)
Canada	1.0	-1.40 (0.072)	-0.308 (0.09)	-0.059 (0.031)	0.068 (0.022)	0	0.22 (0.64)
Australia	1.0	-1.534 (0.051)	0	-0.037 (0.02)	-0.045 (0.022)	0.149 (0.065)	0.54 (0.46)
Sweden	1.0	0	-1.0	-0.098 (0.04)	0	0.095 (0.025)	0.83 (0.84)
Norway	1.0	-1.455 (0.079)	-0.823 (0.09)	-0.163 (0.026)	0	0	1.70 (0.43)
Finland	1.0	0	-1.521 (0.22))	-0.047 (0.02)	0	0.048 (0.023)	2.11 (0.35)
Netherlands	1.0	-0.527 (0.084)	-0.846 (0.056)	-0.141 (0.077)	0	0.327 (0.112)	1.55 (0.21)
Belgium	1.0	0	-1.0	-0.09 (0.033)	0	0.09 (0.034)	2.18 (0.54)
Ireland	1.0	-1.0	-1.0	-0.247 (0.074)	0.081 (0.044)	0	5.67 (0.13)
Hong Kong	1.0	-1.325 (0.058)	-0.138 (0.039)	-0.281 (0.056)	0	0	3.46 (0.17)
Singapore	1.0	-0.788 (0.066)	-0.294 (0.062)	-0.145 (0.028)	0	0	1.86 (0.39)

Note: The table reports the estimated long-run and loading coefficients with asymptotic standard errors in parentheses. ‘Restriction’ reports the likelihood ratio test statistic of the imposed overidentifying restrictions with probability values in parentheses.

The results suggest that in most countries real GDP and real property prices both enter the cointegrating vector. The long-run coefficient of real GDP is not significantly different from zero in Sweden, Finland and Belgium, while the restriction that bank lending is homogenous to real GDP is not rejected for Japan and Ireland. The long-run GDP elasticity is found to be significantly larger than one in Germany, Italy, the UK, Canada, Australia, Norway and Hong Kong and significantly smaller than one in the USA, France, the Netherlands and Singapore.

A zero restriction on the long-run property price coefficient is not rejected for Germany, Italy and Australia, while long-run homogeneity of bank lending to property prices is not rejected for the USA, France, Sweden, Belgium and Ireland. In all other countries, the long-run property price coefficient is significantly smaller than one except for Finland, which is the only case where it is significantly larger than one.

The loading coefficient in the bank lending equation is significantly smaller than zero in every country. The loading coefficient in the GDP equation is significant only in Germany, the UK, Canada, Australia and Ireland, that in the property price equation only in Australia, Sweden, Finland, the Netherlands and Belgium. This finding suggests that it is bank lending rather than real GDP or property prices that takes the system back to long-run equilibrium.

5. Dynamic Interaction

In this section I estimate, based on the VECMs estimated in the previous section, error-correction models for credit growth and the change in property prices. The error-correction models are of the form:

$$(2)\quad \Delta l_t = \gamma_0 CI_{t-1} + \sum_{i=0}^{n} \gamma_{1i} \Delta l_{t-i} + \sum_{j=0}^{n} \gamma_{2j} \Delta y_{t-j} + \sum_{k=1}^{n} \gamma_{3k} \Delta p_{t-k} + \sum_{l=1}^{5} \gamma_{4l} r_{t-l} + \varepsilon_t$$

(3) $\Delta p_t = \lambda_0 CI_{t-1} + \sum_{i=0}^{n} \lambda_{1i} \Delta p_{t-i} + \sum_{j=0}^{n} \lambda_{2j} \Delta y_{t-j} + \sum_{k=1}^{n} \lambda_{3k} \Delta l_{t-k} + \sum_{l=1}^{5} \lambda_{4l} r_{t-l} + \upsilon_t$,

where Δl is real lending growth, Δy is real GDP growth, Δp is the change in real property prices and r is the short-term real interest rate. *CI* is the cointegrating vector linking the levels of real bank lending, real GDP and real property prices reported in Table 2. In each equation I chose the lag order selected for the VECMs in the previous section, which are reported in Table 1, as the maximum lag length n considered for the change in real lending, real GDP growth and the change in real property prices. For the short-term real interest rate I consider up to five lags. The error-correction models were then estimated by applying a general-to-specific modelling strategy eliminating sequentially the least significant lag until all retained lags were significant at least at the 10% level. The full results are reported in Appendix-Tables 1 and 2. In Table 3 and 4 I report the error-correction coefficient and the sum of the coefficients of the retained lags of each variable with t-statistics in parentheses.

Table 3 reports the results for the error-correction model for credit growth. Coefficients, which are significant at least at the 10% level, are in bold. A zero indicates that no single lag was found to be significant so that all lags of the variable were eliminated from the model. The error-correction coefficient is significant in every country, providing strong cross country evidence of long-run causality going from economic activity and property prices to bank lending. There is also evidence of short-run causality going from GDP and property prices to bank lending. In eight countries I find a significant effect of lagged real GDP growth on credit growth. In six countries the effect is significantly positive, while it is significantly negative in Japan and Hong Kong. In thirteen countries I find evidence of short-run causality going from property prices to lending. The overall effect is insignificant in four countries and significantly positive in nine. A significantly negative real interest rate effect is found in only four countries.

Table 4 reports the results for the error-correction models for the change in property prices. I find a significant error-correction coefficient only for four countries, suggesting that on the whole there is rather weak evidence of long-run causality going from bank lending to property prices. There is some evidence of short-run causality going from credit to property prices. In ten countries I find that the sum of the coefficients of lagged credit growth enters significantly the property price error-correction equation and nine countries the total effect is significantly positive. Lags of GDP growth are found to have a significant effect on property prices in five countries and the results are again not clear cut. In the US and Japan the total effect is significantly negative, while it is significantly positive in Italy, Belgium and Singapore.

As regards the effect of real interest rates, the estimation results suggest that real interest rates rather affect property prices than credit. Lags of the short-term real interest rate are found to have a significantly negative effect on property prices in nine countries and on lending in only four.

In order to allow for contemporaneous effects of property prices on bank lending and vice versa, I re-estimate the error-correction models also allowing the current change in property prices to enter the bank lending equations and the current change in bank lending to enter the property price equation. In order to avoid simultaneity bias to affect the estimation, I instrument for the contemporaneous change in property prices and bank lending with the predetermined variables entering the respective error-correction equation. The estimated coefficients of the contemporaneous variables are reported in the last columns of Table 3 and 4. The results suggest that there is only weak evidence of contemporaneous effects in either direction. Only for the US the contemporaneous change in property prices is found to have a significant effect on credit growth.

Table 3: Short-run bank lending dynamics

	CI_{t-1}	Δy_{t-j}	Δp_{t-k}	r_{t-l}	$\bar{R}^2$	Δp_t *(IV)*
USA	**-0.086** **(-3.48)**	**0.293** **(1.78)**	**0.235** **(1.63)**	0	0.72	**0.419** **(2.21)**
Japan	**-0.889** **(-2.47)**	**-0.249** **(-2.00)**	**0.332** **(2.35)**	0	0.62	0.234 (0.75)
Germany	**-0.021** **(-1.73)**	0	-0.009 (-0.16)	**0**	0.29	-0.009 (-0.03)
France	**-0.14** **(-3.63)**	**0.536** **(2.17)**	**0.333** **(2.88)**	**-0.539** **(-2.32)**	0.58	0.298 (1.49)
Italy	**-0.131** **(-5.37)**	0	-0.067 (-0.43)	**-0.121** **(-2.15)**	0.56	0.112 (0.89)
UK	**-0.076** **(-3.55)**	0	**0.092** **(1.63)**	0	0.50	0.241 (1.53)
Canada	**-0.062** **(-2.62)**	0	0	**-0.512** **(-3.12)**	0.38	-0.031 (-0.25)
Australia	**-0.064** **(-3.66)**	**0.30** **(2.06)**	0	0	0.63	0.311 (0.81)
Sweden	**-0.068** **(-2.82)**	0	**0.387** **(3.34)**	0	0.47	-0.317 (-1.08)
Norway	**-0.139** **(-7.14)**	**1.028** **(3.43)**	-0.072 (-0.69)	0	0.60	0.011 (0.04)
Finland	**-0.0464** **(-4.48)**	0	**0.105** **(1.82)**	0	0.58	-0.34 (-0.80)
Neherlands	**-0.108** **(-2.56)**	**0.263** **(1.63)**	-0.063 (-0.70)	0	0.57	0.101 (0.32)
Belgium	**-0.075** **(-3.14)**	**0.433** **(2.57)**	**0.234** **(2.06)**	0.332 (1.19)	0.40	0.128 (0.34)
Ireland	**-0.158** **(-2.98)**	0	**0.628** **(3.53)**	**-0.627** **(-2.57)**	0.69	0.193 (1.20)
Hong Kong	**-0.22** **(-6.03)**	**-0.605** **(-3.93)**	**0.198** **(5.93)**	0	0.55	-0.076 (-0.57)
Singapore	**-0.176** **(-9.23)**	0	0	0	0.55	0.04 (1.06)

Table 4: Short-run property price dynamics

	CI_{t-1}	Δy_{t-j}	Δl_{t-k}	r_{t-l}	$\bar{R}^2$	Δp_t *(IV)*
USA	0	**-0.283 (-2.23)**	**0.213 (2.77)**	0	0.59	0.092 (0.68)
Japan	0	**-0.18 (-2.03)**	**0.321 (4.87)**	**-0.417 (-2.96)**	0.78	0.077 (0.55)
Germany	0	0	0	**-0.728 (-2.51)**	0.69	-0.214 (-0.69)
France	0	0	0	**-0.294 (-1.85)**	0.74	0.004 (0.03)
Italy	0	**1.494 (3.87)**	**0.382 (2.53)**	0	0.73	0.049 (0.27)
UK	0	**0**	**0.477 (2.72)**	0	0.30	-0.171 (-0.31)
Canada	0	0.055 (0.12)	0	**-1.051 (-2.09)**	0.13	0.745 (1.28)
Australia	**0.172 (3.26)**	0	**0.737 (3.05)**	**-0.79 (-2.74)**	0.31	0.119 (1.43
Sweden	0	0	0	**-0.425 (-2.69)**	0.66	-0.037 (-0.27)
Norway	0	0	**0.3586 (2.87)**	**-0.833 (-1.62)**	0.37	0.064 (0.25)
Finland	**0.049 (4.40)**	0	**0.591 (3.89)**	**-0.781 (-2.43)**	0.73	0.732 (0.44)
Netherlands	**0.13 (2.22)**	0	**0.351 (2.24)**	0	0.42	0.241 (0.73)
Belgium	**0.095 (4.05)**	**0.225 (2.04)**	**0.236 (2.38)**	0	0.24	0.213 (1.50)
Ireland	0	0	-0.042 (-0.27	0	0.36	0.174 (0.95)
Hong Kong	0	0	0	0	0.43	0.165 (0.39)
Singapore	0	**0.649 (2.44)**	0	**-2.85 (-2.46)**	0.51	-0.357 (-0.94)

5. Conclusions

Over the last two decades, bank lending has been closely correlated with property prices in both industrialised and developing countries. Theory suggests that property prices may affect credit via various wealth effects, while credit may affect property prices via various liquidity effects. From previous empirical studies it is not clear which effect dominates, since the focus is usually on either effect but not on both. In this study I analyse the direction of causality between bank lending and property prices in 16 industrialised countries over the last two decades.

Cointegration analysis suggests that in all countries there is evidence of a long-run relationship between bank lending, economic activity and property prices, with long-run causality going from property prices and economic activity to bank lending. This finding suggests that property price cycles, reflecting changing beliefs about future economic prospects, drive credit cycles, rather than excessive bank lending, in the wake of financial liberalisation, being the cause of property price bubbles.

The results of estimating error-correction equations for credit growth and the change in property prices suggest that there is evidence of short-run causality going in both directions. This implies that a mutually re-enforcing element in past boom bust cycles in credit and property markets cannot be ruled out.

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