Bank Lending and Property Prices: Some International Evidence*

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ABSTRACT

This paper analyses the patterns of dynamic interaction between bank lending and property prices based on a sample of 20 countries using both time series and panel data techniques. 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. There is also evidence of short-run causality going in both directions, implying that a mutually re-enforcing element in past boom bust cycles in credit and property markets cannot be ruled out.

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1. Introduction

Over the last couple of years, the coincidence of cycles in credit and property markets has been widely documented and discussed in the policy oriented literature (IMF, 2000, BIS, 2001). However, the question of the direction of causality between bank lending and property prices has remained a rather unexplored issue. From a theoretical point of view, causality may go in both directions. Property prices may affect bank lending via various wealth effects. First, due to financial market imperfections, households and firms may be borrowing constrained. As a result, households and firms can only borrow when they 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 therefore an important determinant of the private sector's borrowing capacity. Second, a change in property prices may have a significant effect on consumers' perceived lifetime wealth[2], inducing them to change their spending and borrowing plans and thus their credit demand in order to smooth consumption over the life cycle[3]. 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 property[4]. Property prices therefore influence the risk taking capacity of banks and thus their willingness to extend loans.

On the other hand, bank lending may affect property prices via various liquidity effects. 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 the availability of credit may lower interest rates and stimulate 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. Property can 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

[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] Data on the composition of household wealth, reported in OECD (2000), show that households hold a large share of their wealth in property.

[3] 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).

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.

This potential two-way causality between bank lending and property prices may give rise to mutually reinforcing cycles in credit and property markets[5]. 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 may 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 enters multivariate probit-logit models of financial crisis in industrialised and developing countries. Borio and Lowe (2002) show that a measure of the aggregate asset price[6] 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.

Borio, Kennedy and Prowse (1994) investigate the relationship between credit to GDP ratios and aggregate asset prices 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

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[4] 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).

[5] The possibility of mutually re-enforcing cycles in credit and asset markets has already been stressed by Kindleberger (1978) and Minsky (1982).

[6] 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.

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. Based on this finding, they conclude that bank lending has contributed significantly to the real estate bubble in Asia prior to the 1997 East Asian crisis.

All these studies potentially suffer from simultaneity problems because the potential two-way relationship between credit and property prices is not controlled for. In two recent studies, Hofmann (2001) and Gerlach and Peng (2002) analyse the relationship between bank lending and property prices based on a multivariate empirical framework. Hofmann (2001) finds for a set of 16 industrialised countries that including property prices in the empirical model is imperative for the explanation of the long-run development of bank lending and that long-run causality goes from property prices and real activity to bank lending. Based on impulse response analysis Hofmann (2001) finds that property price innovations have a significantly positive effect on bank lending and vice versa, suggesting a two-way relationship between credit and property prices. The problem with this paper's analysis is that the identified patterns of causality are likely not to be invariant to the indentifying assumptions imposed upon the estimated VARs. Gerlach and Peng (2002) overcome this problem by analysing the direction of causality between bank lending and property prices in Hong Kong based on standard regression techniques, controlling for potential simultaneity problems. They find that long-run and short-run causality goes from property prices to lending, rather than conversely.

In the following I assess, based on time series and panel data techniques, the patterns of dynamic interaction between bank lending and property prices for a sample of 20 countries since the mid 1980s. The plan of the paper is as follows. The following Section 2 describes the data. Section 3 tests and estimates long-run relationships between bank lending, economic activity and property prices. In Section 4 I estimate error-correction models and test for the patterns of long-run and short-run causality between bank lending and property prices. Section 5 concludes.

2. Data

In the following sections I analyse the relationship between real aggregate bank lending, real GDP as a measure of aggregate economic activity, real residential property prices and real money market interest rates in 20 countries: the US, Japan, Germany, France, Italy, the UK, Canada, Switzerland, Sweden, Norway, Finland, Denmark, Spain, the Netherlands, Belgium, Ireland, Australia, New Zealand, Hong Kong and Singapore. The data for the industrialised countries were taken from the IMF and 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.

Bank lending, which was transformed into real terms by deflation with the consumer price index (CPI), is defined as total credit to the private non-bank sector. Cross-country comparisons of the development of bank lending are flawed by differences in the definition of total credit across countries. These differences in definition will be reflected in the results of the empirical analysis. Differences exist, for example, with respect to the treatment of non-performing loans (NPLs) in national credit aggregates. A drop in property prices will on the one hand have a negative effect on the extension of new loans. On the other hand it will give rise to an increase in NPLs. The estimated effect of property prices on bank lending will therefore depend on whether banks are forced to write off NPLs quickly or not. For instance, Japan and the Nordic countries experienced severe banking crises in the late 1980s or early 1990s, which were preceded by a collapse in property prices[7]. While NPLs were quite quickly cleansed from banks' balance sheets in the Nordic countries, this was not the case in Japan[8].

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

[7] Drees and Pazarbasioglu (1998) provide a survey on the causes and consequences of the banking crises in the Nordic countries. The literature on the Japanese crisis is of course enormous. See Hoshi and Kashyap (1999) for a recent survey and the references therein.
[8] For a more detailed discussion of this issue see BIS (2001b).

obtain a measure of real property prices, nominal property prices were deflated with the CPI. Residential property prices may not fully capture the property price developments which are relevant for aggregate bank lending. Credit aggregates comprise bank lending to households and enterprises. The appropriate measure of property prices for the empirical analysis would therefore be an aggregate property price index, comprising both residential and commercial property prices. In Hofmann (2001) I construct such an index as a weighted average of residential and commercial property prices. For most countries, the available commercial property price data are available only in annual frequency and represent only price developments in the largest urban area of the country. The use of these data in empirical analysis is therefore quite problematic. In the few countries where high quality commercial property price data are available, such as Japan, Hong Kong and Singapore, residential and commercial property prices are closely correlated, suggesting that residential property prices may act as a proxy for omitted commercial property prices in the empirical analysis.

The short-term real interest rate is measured as the three months interbank money market rate less four quarter CPI inflation. The short-term real money market rate serves as a proxy for real aggregate financing costs. A more accurate measure would be an aggregate lending rate. Representative lending rates are, however, not available for most countries. Empirical evidence suggests that lending rates are tied to money market rates[9], implying that money market rates are a valid approximation of financing costs.

3. Unit Roots, Cointegration and Long-run Relationships

The sample period for the following analysis is 1985:1 – 2001:4. This is a rather short-sample period, implying that the power of unit root and cointegration tests may be low. The results of these tests should therefore be taken with caution and be regarded as being rather suggestive. As a tentative attempt to partly overcome this problem I exploit the rather large cross-section dimension of my analysis to perform panel unit root and cointegration tests.

As a first step I perform standard augmented Dickey-Fuller (ADF) unit root tests (Dickey and Fuller, 1981) to test for the order of integratedness of the time series under investigation. The ADF test regression is of the form:

$$(1)\quad \Delta x = \mu + \delta t + \gamma x_{t-1} + \sum_{i=1}^{k} \Delta x_{t-i} + \varepsilon_t \ .$$

Allowing for a maximum lag order of four, the lag order k was determined by sequential t-tests eliminating all lags up to the first significant at the 5% level. The test regression for the level of each variable contained a constant and a trend, the test regression for the first difference only a constant. The ADF test statistic is the t-statistic of γ. If γ is significantly smaller than zero, the null hypothesis of a unit root can be rejected. The 1%, 5% and 10% critical values are respectively -4.10, -3.48 and -3.17 for the level tests and -3.53, -2.90 and -2.59 for the tests for the first differences[10]. The results are displayed in Table 1. In the last row of Table 1 I also report a panel ADF test proposed by Im et al. (2003). They show that the standardised average of the N individual ADF test statistics

$$(2)\quad z_{ADF} = \frac{\sqrt{N}(\bar{t}_\gamma - \mu)}{\sqrt{\delta}}$$

has a standard normal distribution, where $\bar{t}_\gamma$ is the average of the individual ADF test statistics and μ and δ are respectively the mean and the variance of the distribution of the ADF test statistic. The appropriate mean and variance adjustment values are tabulated in Im et al. (2003). The test is one sided. The 1%, 5% and 10% critical values are -1.96, -1.64 and -1.28. Large negative values therefore imply a rejection of the null of a unit root.

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[9] See Borio and Fritz (1995) for a large sample of industrialised countries, Hofmann (2002) for euro area countries and Hofmann and Mizen (2002, 2003) for the UK.
[10] Critical values were calculated based on the response surfaces reported in MacKinnon (1991).

Table 1: Unit root tests

	Real Lending		Real GDP		Real house prices		Real interest rates	
	Level	1st difference	Level	1st difference	Level	1st difference	Level	1st difference
Australia	-2.48	-2.17	-1.51	-5.98***	-2.65	-3.86***	-4.01**	-4.31***
Belgium	-1.05	-3.78***	-2.85	-8.31***	-2.21	-4.43***	-2.01	-7.98***
Canada	-2.31	-4.49***	-1.73	-4.83***	-2.55	-2.96**	-2.70	-6.17***
Switzerland	-1.43	-1.87	-2.31	-9.64***	-2.75	-2.55	-3.70**	-6.16***
Denmark	-1.42	-4.76***	-1.89	-5.63***	-1.51	-2.53	-2.02	-7.72***
Spain	-2.63	-2.06	-1.78	-2.59*	-1.85	-3.65***	-3.48*	-6.38***
Finland	-2.45	-1.26	-2.15	-2.32	-2.68	-2.48	-2.02	-7.72***
France	-2.90	-1.49	-2.76	-5.48***	-3.20*	-2.29	-2.51	-8.07***
Germany	-2.32	-5.75***	-1.59	-2.79*	-1.88	-2.67*	-2.19	-5.92***
Hong Kong	-0.87	-4.77***	-2.66	-4.41***	-0.15	-4.65***	-1.74	-7.96***
Ireland	-1.26	-2.61*	-2.44	-2.05	-1.93	-2.50	-2.54	-8.75***
Italy	-1.91	-3.36**	-2.25	-7.78***	-3.08	-2.13	-2.94	-7.06***
Japan	-1.52	-1.22	-0.94	-2.19	-2.16	-1.74	-2.54	-9.70***
Netherlands	-1.52	-1.99	-1.88	-2.30	-2.09	-3.15**	-2.28	-5.23***
Norway	-2.21	-2.47	-1.75	-3.50**	-1.42	-2.86*	-3.32*	-6.66***
New Zealand	-1.60	-1.91	-1.36	-8.39***	-2.03	-5.46***	-4.71***	-11.76***
Singapore	-2.45	-2.68*	-1.75	-4.44***	-1.70	-3.78***	-2.75	-5.59***
Sweden	-1.85	-4.13***	-1.73	-3.30**	-2.54	-2.98**	-4.83***	-6.41***
UK	-1.69	-3.94***	-2.77	-2.52	-3.45*	-2.32	-3.84**	-6.71***
US	-2.24	-2.26	-2.10	-3.50**	-1.30	-1.82	-1.81	-4.79***
Panel	1.86	-7.58***	1.66	-16.21***	0.49	-7.85***	-4.82***	-29.33***

*Note: The table reports ADF tests statistics for the null hypothesis of a unit root. The 1%, 5% and 10% critical values are respectively -4.10, -3.48 and -3.17 for the level tests and -3.53, -2.90 and -2.59 for the test for the first differences. The 1%, 5% and 10% critical values for the panel tests are -1.96, -1.64 and -1.28. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.*

On the whole, the results suggest that the natural logs of real bank lending, real property prices and real GDP are integrated of order one. This conclusion is suggested both by the individual country level tests as well as by the panel tests. The short-term real interest rate appears to be a borderline case. The null of non-stationarity is rejected at least at the 10% level in seven countries out of 13 countries. The panel unit root test strongly suggests that the real interest rate is a stationary process.

Given the results of the unit root tests we test in the following for the presence of a long-run relationship between real bank lending, real GDP and real property prices. The level of the real interest rate is not allowed to enter the long-run relationship. Since we have more than two variables in the system and therefore potentially more than one long-run relationship among these variables, the multivariate Johansen approach (Johansen, 1988, 1991, 1992), which allows to test for the number of cointegrating relationships in the system, is a natural starting point for cointegration analysis. The Johansen approach is based on maximum likelihood estimation of a cointegrating VAR model, which can be formulated in vector error correction form:

(3) $$\Delta x_t = \Gamma_1 \Delta x_{t-1} + ... + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \varepsilon_t .$$

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 μ unrestricted. The rank of the matrix Π indicates the number of long-run relationships between the endogenous variables in the system[11]. The cointegrating rank hypothesis for the Johansen trace test is specified as $H(r): rank(\Pi) \leq r$ against the alternative $H(p): rank(\Pi) = p$.

The lag order of the VECMs was determined based on sequential Likelihood-ratio tests, eliminating all lags up to the first lag significant at the 5% level. The results of the trace test

[11] For a detailed technical exposition of the Johansen approach see Johansen (1995).

are reported in Table 2. The 1%, 5% and 10% critical values are respectively 35.65, 29.68 and 26.79 for $H(0): rank(\Pi) = 0$, 20.04, 15.41 and 13.33 for $H(0): rank(\Pi) \leq 1$ and 6.65, 3.76 and 2.69 for $H(0): rank(\Pi) \leq 2$ [12]. I also report the result of a panel cointegration trace test proposed by Larsson et al. (2001). The test statistic is the standardised average of the N individual trace test statistics:

$$(4)\ \Psi_{LR} = \frac{\sqrt{N}(\overline{LR} - \mu)}{\sqrt{\delta}},$$

where $\overline{LR}$ is the average of the individual trace test statistics and μ and δ are respectively the mean and the variance of the asymptotic distribution of the trace test statistic, which are tabulated in Osterwald-Lennum (1992). Larsson et al. (2001) show that the test statistic has a standard normal distribution. The test is one sided with large positive values of the test statistics suggesting a rejection of the null hypothesis of no cointgration. The 1%, 5% and 10% critical values are therefore 1.96, 1.64 and 1.28 respectively.

The results suggest that there is a single long-run relationship between bank lending, GDP and property prices. The null of no cointegration is rejected at least at the 10% level in 15 out of 20 countries. Two long-run relationships are indicated only for the UK. The panel trace test rejects the null hypothesis of one long-run relationship against the alternative of no long-run relationship, but it does not reject the null hypothesis of one long-run relationship against the alternative of more than one, therefore suggesting that there is a single long-run relationship in the system.

In order to cross-check the results of the Johansen test I also report in Table 2 the results of two alternative residual based cointegration tests. Residual based cointegration tests were first discussed in Engle and Granger (1987) and are based on tests of a unit root hypothesis in the residuals $\hat{\varepsilon}_t$ of the estimated cointegrating relationship

$$(6)\ \hat{y}_t = \hat{\alpha} + \hat{\beta}_1 x_{1t} + \hat{\beta}_2 x_{2t} + \hat{\varepsilon}_t\ ,$$

[12] Critical values were taken from Osterwald-Lennum (1992).

where in our case y is the log of real bank lending, x_1 is the log of real GDP and x_2 is the log of real property prices. I calculate the long-run residuals $\hat{\varepsilon}_t$ based on the dynamic OLS (DOLS) estimator of the long-run coefficients proposed by Saikkonen (1991) and Stock and Watson (1993), which controls for regressor endogeneity and serial correlation in the long-run regression by adding leads and lags of the first difference of the regressors to the estimating equation. I then apply a standard ADF unit root test and a Lagrange multiplier (LM) test proposed by Harris and Inder (1994). The former tests the null of no cointegration, while the latter takes cointegration as the null.

The ADF cointegration test regression is given by:

$$(6)\ \Delta\hat{\varepsilon}_t = \mu + \gamma\hat{\varepsilon}_{t-1} + \sum_{i=1}^{k}\Delta\hat{\varepsilon}_{t-i} + \varsigma_t$$

The lag order k was determined by sequential t-tests eliminating all lags up to the first significant at the 5% level, allowing for a maximum lag order of four. The 1%, 5% and 10% critical values were calculated based on the response surfaces reported in MacKinnon (1991) and are respectively -4.50, -3.86 and -3.54. I also report a panel ADF cointegration test proposed by Pedroni (1999). The test statistic is the standardised average of the N individual ADF cointegration test statistics:

$$(7)\ \Psi_{ADF} = \frac{\sqrt{N}(\bar{t}_\gamma - \mu)}{\sqrt{\delta}},$$

which has again a standard normal distribution. The test is again one sided with large negative values suggesting a rejection of the null. The results, which are reported in the second-last column of Table 2, clearly suggest that the null of no cointegration cannot be rejected. Not a single test statistic is statistically significant.

The LM test proposed by Harris and Inder (1994) is a multivariate extension of the unit root test proposed by Kwiatkowski et al. (1992). The test statistic is given by:

$$(8)\quad LM = \frac{\sum_{t=1}^{T} S_t^2}{\hat{\sigma}_\varepsilon^2}$$

where $S_t^2 = \sum_{i=1}^{t} \hat{\varepsilon}_i$ is the running partial sum of the residuals and $\hat{\sigma}_\varepsilon$ is the estimated residual variance. The 1%, 5% and 10% critical values are respectively 0.37, 0.22 and 0.16 (Harris and Inder, 1994). McCoskey and Kao (1998) proposed a panel version of the Harris and Inder cointegration test. The test statistic is the standardised average of the *N* individual LM test statistics:

$$(9)\quad \Psi_{LM} = \frac{\sqrt{N}(\overline{LM} - \mu)}{\sqrt{\delta}}.$$

This test statistic is again one sided with a standard normal distribution. Large positive values suggest a rejection of the null. The test results are reported in the last column of Table 2 and suggest that the null of cointegration cannot be rejected for about half of the countries and also not by the panel test.

What do we make of these results? The Johansen and the Harris and Inder test suggest that there exists a (single) long-run relationship linking bank lending, GDP and property prices, while the ADF cointegration test suggests that there is no cointegration. Monte Carlo evidence on the performance of cointegration tests in small samples, which is summarised in Maddala and Kim (1998), appears to be inconclusive on which test is more reliable. A priori it is therefore unclear which test we should trust more. The cointegration test results do not provide compelling evidence in favour of cointegration, but neither do they provide compelling evidence against it. With a view to the following section, the main conclusion of this section is therefore that cointegration tests do not make a clear case against the inclusion of a long-run relationship between bank lending and property prices in dynamic models of bank lending and property price interaction.

Table 2: Cointegration Test Results

	Johansen Test			*Residual Based Tests*	
	$r = 0$	$r \leq 1$	$r \leq 2$	*ADF test*	*LM test*
Australia	30.48**	9.37	0.74	-2.55	0.14
Belgium	28.5*	8.64	0.42	-1.31	0.19
Canada	27.78*	8.33	0.05	-1.68	0.24**
Switzerland	27.45*	9.92	1.7	-0.74	0.20*
Denmark	22.94	8.26	0.84	-1.78	0.22**
Spain	35.72***	8.29	0.49	-1.11	0.15
Finland	54.24***	13.29	0.48	-1.81	0.24**
France	32.45**	10.03	0.68	-0.78	0.22**
Germany	30.58**	8.73	1.12	-2.52	0.24**
Hong Kong	21.13	6.31	0.01	-3.44	0.07
Ireland	25.41	4.76	0.56	-2.56	0.13
Italy	19.22	4.24	0.07	-1.56	0.11
Japan	53.36***	11.19	1.75	-3.47	0.08
Netherlands	20.88	5.1	0.14	-2.88	0.18*
Norway	48.58***	11.22	5.27	-1.72	0.2
New Zealand	26.96*	9.63	2.99	-2.78	0.17
Singapore	28.48*	3.82	0.1	-3.35	0.11
Sweden	31.05**	13.37	1.92	-2.58	0.24**
UK	38.41***	18.70**	2.95	-1.51	0.23**
US	29.49*	10.75	0.34	-2.18	0.19*
Panel	9.69***	1.14	0.44	-0.70	1.02

Note: 'Johansen test' displays the test statistics of the Johansen trace test for cointegration. The 1%, 5% and 10% critical values for the cointegration test are 35.65, 29.68 and 26.79 for r=0, 20.04, 15.41 and 13.33 for $r \leq 1$ and 6.65, 3.76 and 2.69 for $r \leq 2$ *(Osterwald-Lennum, 1992). The 1%, 5% and 10% critical values for the panel cointegration test are respectively 1.96, 1.64 and 1.28. 'ADF test' reports the ADF cointegration test statistic. The 10% critical value is -3.54 for the individual country tests and -1.28 for the panel test. 'LM test' reports the results of the Harris and Inder cointegration test. The 1%, 5% and 10% critical values are 0.37, 0.22 and 0.16 for the individual country tests. The 10% critical value for the panel test is 1.28. *, ** and *** indicates significance of a test statistic at the 10%, 5% and 1% level respectively.*

As the next step we proceed to estimate the long-run coefficients linking real bank lending, real GDP and real property prices. Since we have at most one long-run relationship there are various approaches that are applicable. In Table 3 we present estimates from the Johansen Maximum Likelihood procedure, which are obtained by imposing identifying restrictions on the cointegrating vector. We also report estimates obtained from the Fully Modified OLS (FM-OLS) approach proposed by Phillips and Hansen (1990) and from the DOLS approach proposed by Saikkonen (1991) and Stock and Watson (1993). The former controls for regressor endogeneity and serial correlation by applying a non-parametric correction to the OLS estimators of the long-run coefficients in (6), while the latter adds leads and lags of the first difference of the regressors to the long-run regression equation. For both the FM-OLS and the DOLS estimator we also report panel estimates based on fixed effects panel estimators proposed respectively by Pedroni (1999) and Kao and Chiang (2000).

For each estimator we report the coefficient estimates with t-statistics in parentheses. The results suggest that in most countries real GDP and real property prices both enter significantly the cointegrating vector. However, while the estimates obtained from the FM-OLS and the DOLS estimators are broadly similar, the estimates obtained from the Johansen procedure are sometimes quite different and counter-intuitive. Monte Carlo studies, summarised in Maddala and Kim (1998), suggest that the Johansen procedure exhibits larger variation than single equation estimators in small samples. The Monte Carlo evidence also suggests that the DOLS approach is preferable to the FM-OLS estimator in small samples. Kao and Chiang (2000) report Monte Carlo evidence suggesting that also the panel DOLS estimator outperforms the panel FM-OLS estimator in small samples. For this reason, the error-correction terms which are included in the error-correction models estimated in the following section were calculated using the DOLS estimates of the long-run coefficients.

Table 3: Long-run Relationships

	Johansen ML		Fully Modified OLS		Dynamic OLS	
	Real GDP	Real house prices	Real GDP	Real house prices	Real GDP	Real house prices
Australia	1.97	-0.78	1.34	0.40	1.41	0.33
	(11.16)	(-2.76)	(12.10)	(2.58)	(12.55)	(1.92)
Belgium	2.67	-2.77	0.24	0.97	0.49	0.85
	(2.61)	(-4.89)	(0.40)	(3.00)	(0.83)	(2.78)
Canada	1.37	0.002	1.42	0.26	1.40	0.28
	(13.38)	(0.02)	(15.40)	(2.13)	(18.11)	(4.41)
Switzerland	4.90	0.12	1.92	0.16	2.06	0.19
	(6.46)	(0.43)	(11.83)	(2.49)	(10.97)	(3.77)
Denmark	0.80	0.47	0.65	0.24	0.75	0.19
	(2.26)	(1.89)	(1.75)	(1.07)	(4.38)	(1.47)
Spain	2.66	-0.56	1.99	-0.02	2.04	-0.10
	(11.18)	(-5.03)	(11.70)	(-0.21)	(12.98)	(-0.94)
Finland	-0.31	1.41	0.02	0.32	0.23	0.23
	(-1.62)	(8.64)	(0.06)	(1.43)	(0.71)	(1.31)
France	0.83	0.92	0.94	0.78	0.95	(0.88
	(15.12)	(13.55)	(8.86)	(5.84)	(26.43)	(19.23)
Germany	2.17	-0.69	2.23	-0.47	2.22	-0.46
	(36.06)	(-9.86)	(26.72)	(-5.67)	(66.39)	(-8.63)
Hong Kong	2.17	0.001	1.27	0.21	1.31	0.18
	(6.18)	(0.01)	(12.96)	(3.17)	(13.34)	(3.06)
Ireland	1.12	0.79	0.95	0.84	0.89	0.88
	(8.74)	(6.29)	(10.35)	(9.05)	(16.41)	(19.04)
Italy	2.04	2.36	2.05	0.24	1.98	0.25
	(1.91)	(3.66)	(14.32)	(3.05)	(26.22)	(4.49)
Japan	2.73	0.47	1.13	0.53	1.20	0.56
	(12.84)	(2.49)	(32.90)	(12.88)	(44.03)	(26.56)
Netherlands	0.26	1.03	0.54	0.87	0.46	0.93
	(1.42)	(10.09)	(3.17)	(10.21)	(3.11)	(12.12)
Norway	1.48	0.77	1.57	0.39	1.59	0.44
	(49.04)	(9.56)	(9.05)	(2.76)	(35.61)	(7.26)
New Zealand	1.32	0.90	2.12	0.05	1.82	0.28
	(2.69)	(2.04)	(7.24)	(0.17)	(6.72)	(1.45)
Singapore	0.83	0.24	0.92	0.14	0.91	0.17
	(16.61)	(5.34)	(8.98)	(1.56)	(16.83)	(3.79)
Sweden	0.74	-0.42	0.42	0.78	0.58	0.78
	(1.44)	(-1.11)	(1.63)	(5.14)	(2.14)	(6.00)
UK	1.82	0.35	1.98	0.26	1.98	0.35
	(13.53)	(3.27)	(9.88)	(1.69)	(19.47)	(2.99)
US	1.49	1.67	1.63	0.97	1.63	1.07
	(25.07)	(6.79)	(16.50)	(3.44)	(38.43)	(6.92)
Panel	-	-	1.27	0.40	1.22	0.26
			(48.25)	(14.71)	(11.15)	(5.12)

4. Dynamic Interaction

In this section I estimate error-correction models (ECMs) for credit growth and the change in property prices in order to investigate the patterns of long-run and short-run causality between bank lending and property prices. The ECMs are of the form:

$$(10)\ \Delta l_t = \gamma_0 CI_{t-1} + \sum_{i=0}^{4} \gamma_{1i} \Delta l_{t-i} + \gamma_2 \Delta y_t + \gamma_3 \Delta p_t + \gamma_4 \Delta r_{t-1} + \varepsilon_t$$

$$(11)\ \Delta p_t = \lambda_0 CI_{t-1} + \sum_{i=0}^{4} \lambda_{1i} \Delta p_{t-i} + \lambda_2 \Delta y_t + \lambda_3 \Delta l_t + \lambda_4 \Delta r_{t-1} + \upsilon_t ,$$

where Δl is real lending growth, Δy is real GDP growth, Δp is the change in real property prices and Δr is the change in the short-term real interest rate. *CI* is the DOLS estimate of the cointegrating vector linking the levels of real bank lending, real GDP and real property prices reported in Table 3. Estimating equation (10) and (11) for 20 countries yields two systems of 20 equations each for the change in real lending and the change in real property prices. In order to prevent simultaneity bias from affecting the estimation the contemporaneous variables included in each equation were instrumented for using four own lags as instruments. The systems were estimated by three stage least squares in order to account for potential contemporaneous correlation in the errors across equations. Tables 4 and 5 report the estimation results. Coefficients are reported with t-statistics in parentheses. The last row reports the results of a pooled fixed effects three-stage least squares regression.

Table 4: Short-run bank lending dynamics

	CI_{t-1}	Δy_t	Δp_t	Δr_{t-1}
Australia	-0.046*	0.255	0.076	0.364***
	(-1.69)	(0.89)	(0.94)	(-4.23)
Belgium	-0.023	0.455	-0.575	0.094
	(-0.54)	(0.93)	(-1.38)	(0.40)
Canda	-0.056**	0.158	0.081	-0.046
	(-2.49)	(0.79)	(1.22)	(-0.45)
Switzerland	0.018	0.724	0.457*	0.162
	(0.29)	(0.90)	(1.78)	(0.77)
Denmark	-0.070*	-0.059	-0.096	-0.522*
	(-1.94)	(-0.11)	(-0.34)	(-1.93)
Spain	-0.076***	-0.501**	0.234***	0.111*
	(-4.62)	(-2.57)	(4.50)	(1.65)
Finland	-0.032*	0.508	-0.193	0.222
	(-1.70)	(1.01)	(-1.35)	(-0.92)
France	-0.190***	1.083**	0.030	0.344**
	(-3.38)	(2.17)	(0.13)	(2.32)
Germany	0.042	-0.034	0.193**	0.152
	(1.52)	(-0.24)	(2.45)	(1.26)
Hong Kong	-0.146***	0.417	0.206***	0.151
	(-4.27)	(1.36)	(3.63)	(1.05)
Ireland	-0.205***	0.303	0.183	0.130
	(-4.05)	(0.97)	(1.03)	(1.35)
Italy	-0.136***	1.014	-0.111	-0.151
	(-3.05)	(0.94)	(-0.87)	(-0.88)
Japan	-0.134**	0.842***	-0.011	0.082
	(-2.47)	(3.80)	(-0.61)	(0.63)
Netherlands	-0.102***	0.458	0.234**	0.133
	(-3.20)	(1.52)	(2.32)	(0.93)
Norway	-0.139***	1.455**	-0.437	-0.209
	(-3.35)	(2.74)	(-1.64)	(-0.88)
New Zealand	-0.061	0.180	0.184	-0.169
	(-1.11)	(0.30)	(0.55)	(-1.60)
Singapore	-0.162***	-0.177	-0.050	-0.377**
	(-6.58)	(-1.43)	(-1.01)	(-2.14)
Sweden	-0.133***	-0.153	0.195	0.047
	(-5.20)	(-0.43)	(1.52)	(0.36)
UK	-0.066***	0.014	0.078	-0.287*
	(-2.75)	(0.03)	(0.82)	(-1.72)
US	-0.012	-0.246	0.160	-0.001
	(-0.15)	(-0.20)	(0.44)	(-0.11)
Panel	-0.027***	0.152***	0.113***	-0.001
	(-7.61)	(2.77)	(6.52)	(-0.23)

Note: The table reports the results from estimating equation 10 jointly for 20 countries by three stage least squares. The change in real GDP and the change in real property prices were instrumented for using respectively four own lags.

Table 5: Short-run property price dynamics

	CI_{t-1}	Δy_t	Δl_t	Δr_{t-1}
Australia	0.151***	-0.425	0.467***	-0.570***
	(3.18)	(-0.62)	(2.77)	(-2.94)
Belgium	0.119***	-0.595	0.577*	-0.397
	(3.55)	(-1.05)	(1.92)	(-1.58)
Canda	-0.014	-0.262	0.269	-0.450**
	(-0.17)	(-0.41)	(0.50)	(-2.01)
Switzerland	0.003	0.186	-0.007	-0.289
	(0.05)	(0.22)	(-0.02)	(-1.26)
Denmark	-0.081***	-0.203	0.168	-0.414*
	(-3.74)	(-0.61)	(1.55)	(-2.45)
Spain	0.016	-0.723	0.244	-0.062
	(0.40)	(-1.21)	(1.22)	(-0.36)
Finland	-0.075**	-1.426**	-0.264	-0.705***
	(-2.20)	(-1.83)	(-1.01)	(-2.76)
France	-0.027	-0.162	0.207	-0.115
	(-0.51)	(-0.54)	(1.00)	(-1.03)
Germany	-0.186***	0.515**	0.211	-0.388***
	(-5.03)	(2.68)	(0.81)	(-2.87)
Hong Kong	0.208*	0.773	0.840	-0.186
	(1.71)	(1.14)	(1.45)	(-0.39)
Ireland	0.011	1.600***	0.378	-0.127
	(0.10)	(4.36)	(0.89)	(-0.90)
Italy	0.025	2.310*	0.620**	0.068
	(0.48)	(1.65)	(2.03)	(0.29)
Japan	0.019	0.427**	0.242*	0.092
	(0.35)	(2.10)	(1.78)	(0.81)
Netherlands	0.234***	0.541	0.532**	-0.538***
	(4.77)	(1.25)	(2.52)	(-2.63)
Norway	-0.022	-0.015	0.576*	-0.514*
	(-0.29)	(-0.03)	(1.77)	(-1.99)
New Zealand	-0.04	0.487	0.018	-0.233**
	(-0.81)	(1.18)	(0.10)	(-2.09)
Singapore	-0.241	1.389**	-0.110	-0.354
	(-1.58)	(2.64)	(-0.21)	(-0.569)
Sweden	-0.007	-0.354	-0.057	-0.246**
	(-0.24)	(-1.28)	(-0.43)	(-2.31)
UK	-0.015	2.349***	0.569	-0.542*
	(-0.25)	(3.00)	(1.56)	(-1.70)
US	0.073***	-0.251	-0.095	0.054
	(4.04)	(-0.85)	(-1.39)	(0.41)
Panel	-0.003	0.475***	0.136***	-0.241***
	(-0.71)	(6.07)	(3.99)	(-6.74)

Note: The table reports the results from estimating equation 11 jointly for 20 countries by three stage least squares. The change in real GDP and the change in real lending were instrumented for using respectively four own lags.

The results suggest that long-run causality goes from property prices to bank lending, rather than conversely. The error-correction term in the ECM for bank lending is significantly negative in 15 out of 20 cases and the panel estimate is significantly negative at the 1% level. In the ECM for property prices, the error-correction term is significantly positive in 5 cases, but significantly negative in three. The pooled estimate is insignificant. Short-run causality appears to go in both directions. The change in real property prices (real bank lending) has a significantly positive effect on bank lending (property prices) in five (six) countries and the pooled estimate is in both ECMs significant at the 1% level. GDP growth and real interest rates appear to matter more for property prices than for bank lending. GDP growth has a significantly positive effect on property prices (bank lending) in 3 (6) countries. The pooled estimate is in both cases significantly positive at the 1% level. The change in real interest rates has a significantly negative effect on property prices (bank lending) in 10 (2) countries. The pooled estimate is significantly negative at the 1% level in the property prices ECM but insignificant in the bank lending ECM[13].

5. Conclusions

Over the last couple of years, the coincidence of cycles in credit and property markets has been widely documented and discussed in the policy oriented literature. In this paper I analyse the causes of this coincidence. From a theoretical point of view, the relationship between bank lending and property prices is multifaceted. Property prices may affect credit via various wealth effects, while credit may affect property prices via various liquidity effects. Previous empirical studies were not able to disentangle the direction of causality, since the focus was usually on either effect, but not on both.

[13] The finding of bank lending being rather irresponsive to interest rates is consistent with the results of the impulse response analysis in Goodhart and Hofmann (2003).

I analyse the patterns of dynamic interaction between bank lending and property prices based on a sample of 20 industrialised countries using both time series and panel data techniques. Long-run causality appears to go from property prices to bank lending, rather than conversely. 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 liberalization or overly loose monetary policy, being the cause of property price bubbles. However, there is also evidence of short-run causality going in both directions, implying that a mutually re-enforcing element in past boom bust cycles in credit and property markets cannot be ruled out.

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