Asset Prices, Financial Conditions, and the Transmission of Monetary Policy

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ABSTRACT

In this paper we assess the role of asset prices as information variables for aggregate demand conditions and in the transmission of monetary policy. A Monetary Conditions Index, a weighted average of the short-term interest rate and the exchange rate, has commonly been used as a composite measure of the stance of monetary policy and aggregate demand conditions. However, other asset prices, property and share prices, also affect aggregate demand. By looking at reduced form coefficient estimates and VAR impulse responses we derive Financial Conditions Indices, a weighted average of the short-term real interest rate, the effective real exchange rate, real property and real share prices, for the G7 countries. We find that house and share prices get a substantial weight in such an index and that the derived Financial Conditions Indices contain useful information about future inflationary pressures.

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1. Introduction

A Monetary Conditions Index (MCI), a weighted average of the short-term interest rate and the exchange rate, has commonly been used, at least in open economies, as a composite measure of the stance of monetary policy. The MCI concept was based on empirical findings that inflationary pressures are determined by excess aggregate demand and that monetary policy mainly affects aggregate demand via its leverage over short-term interest rates and the real exchange rate. Changes in the stance of monetary policy affect short-term money market interest rates, which in turn influence the investment and saving decisions of households and firms and thus domestic demand conditions. A change in short-term interest rates changes, ceteris paribus, the interest rate differential vis-à vis the rest of the world and may thus lead to a change in the real exchange rate, which in turn affects the competitiveness of domestic firms vis-à-vis foreign firms and thus external demand conditions.

Recent developments in theoretical and empirical research on the monetary transmission process imply that property and equity prices may also play an important role in the transmission of monetary policy via wealth and balance sheet effects. Monetary policy can affect property and equity prices via arbitrage effects and/or a change in discounted expected future dividends, which gives rise to the wealth effect and the balance sheet effect of monetary policy.

A wealth effect may arise because a change in asset prices affects the financial wealth of consumers, which may induce them to change their consumption plans (Modigliani, 1971). The strength of the wealth effect of a change in asset prices depends in part on the share of the respective asset in private sector wealth. Table 1 shows the composition of household wealth in the G7 countries in 1998. In all countries, maybe except for Japan, housing assets account for a substantial share in household wealth. Equity accounts for a lower share than housing assets throughout, but there are significant differences across countries. The share of equity in household wealth is considerable in the US, Italy, the UK and Canada, but negligible in Japan, Germany and France.

Table 1: Composition of household total assets (in percent)

	Housing assets	Equity	Other financial assets	Other tangible assets
United States	21	20	50	8
Japan	10	3	44	43
Germany	32	3	35	30
France	40	3	47	9
Italy	31	17	39	13
United Kingdom	34	12	47	7
Canada	21	17	39	23

Note: Data refer to 1998 (1997 for France)
Source: OECD Economic Outlook, December 2000, Table VI.1

In the appendix we also show the evolution of the share of housing assets in household total assets and of the share of equity in net financial wealth. While the share of housing wealth in total wealth has been rather stable since 1970, the share of equity in net financial wealth has been rising in most countries.

The balance-sheet channel occurs in part as a result of problems arising from asymmetric information in the credit market, which gives rise to adverse selection and moral hazard concerns. The lower the net worth of firms and households, the more severe these problems will be, since there will be less collateral available to secure loans. A rise in asset prices raises the borrowing capacity of firms and households by increasing the value of collateral. The additionally available credit can be used to purchase goods and services and thus stimulates economic activity (Bernanke and Gertler, 1989, Kiyotaki and Moore, 1997 and Bernanke, Gertler and Gilchrist, 1998). Table 2 shows the share of real estate secured borrowing in the G7 countries. The share of property secured borrowing lies around 60% in the Anglo-Saxon countries and around 40% in the Continental European countries. For Japan no estimate is available, but there are indications suggesting that the share might be quite high (Borio, 1996).

Table 2: Share of loans secured by real estate collateral (in percent)

United States	Germany	France	Italy	United Kingdom	Canada
66	36	41	40	59	56

Note: Data refer to 1993
Source: Borio(1996), Table 12, p.101.

Thus, from a theoretical point of view there is a strong case to consider other asset prices besides interest and exchange rates as indicators of the stance of monetary policy as well. In the following we will try to derive an extended Monetary Conditions Index or Financial Conditions Index (FCI) for the G7 countries. FCI weights are derived based on coefficient estimates of reduced form demand equations and impulse responses of an identified VAR.

2. Constructing Financial Conditions Indices: Strategies and problems

The general strategy

We are aiming at the construction of an aggregate measure of monetary or financial conditions, a Financial Conditions Index (FCI), consisting of a short-term interest rate, the real effective exchange rate, real house prices and real share prices[1]. Our FCI is defined as $FCI_t = \sum_i w_i(q_{it} - \overline{q}_{it})$, where q_{it} is the price of asset i in period t, $\overline{q}_{it}$ is the long-run trend or equilibrium value of the price of asset i in period t, and w_i is the relative weight given to the price of asset i in the FCI[2].

[1] It would, of course, also have been desirable to include a long-term interest rate, since in many countries long-term interest rates are more relevant for aggregate demand conditions than short-term interest rate. The problem we faced was that we wanted to concentrate on real variables, and it is not clear how to construct a measure of real long-term interest rates. What one would need is a measure of expected inflation for a horizon equal to the maturity of the bond underlying the long-term yield. We tried to construct a measure of inflation expectations based on a three-year moving average of inflation, but the thus constructed real long-term interest rate always appeared with the wrong sign in the output gap equations, indicating that this measure is probably a rather poor proxy of real long-term rates. For this reason we decided to exclude long-term interest rates from the analysis. But we are planning to address this issue in future work.

[2] This definition does not conform to the usual definition of a MCI, where the interest rate and the exchange rate are set in relation to their level in a base period. Since house prices and share prices are trending variables, such a definition would have led to trending FCIs. To avoid that we set our asset prices in relation to some measure of long-run trend or equilibrium.

As with the construction of MCIs, there are three possibilities for estimating the weights of the respective asset prices in an FCI:

- simulations in large scale macro-econometric models
- reduced form aggregate demand equations
- VAR impulse responses.

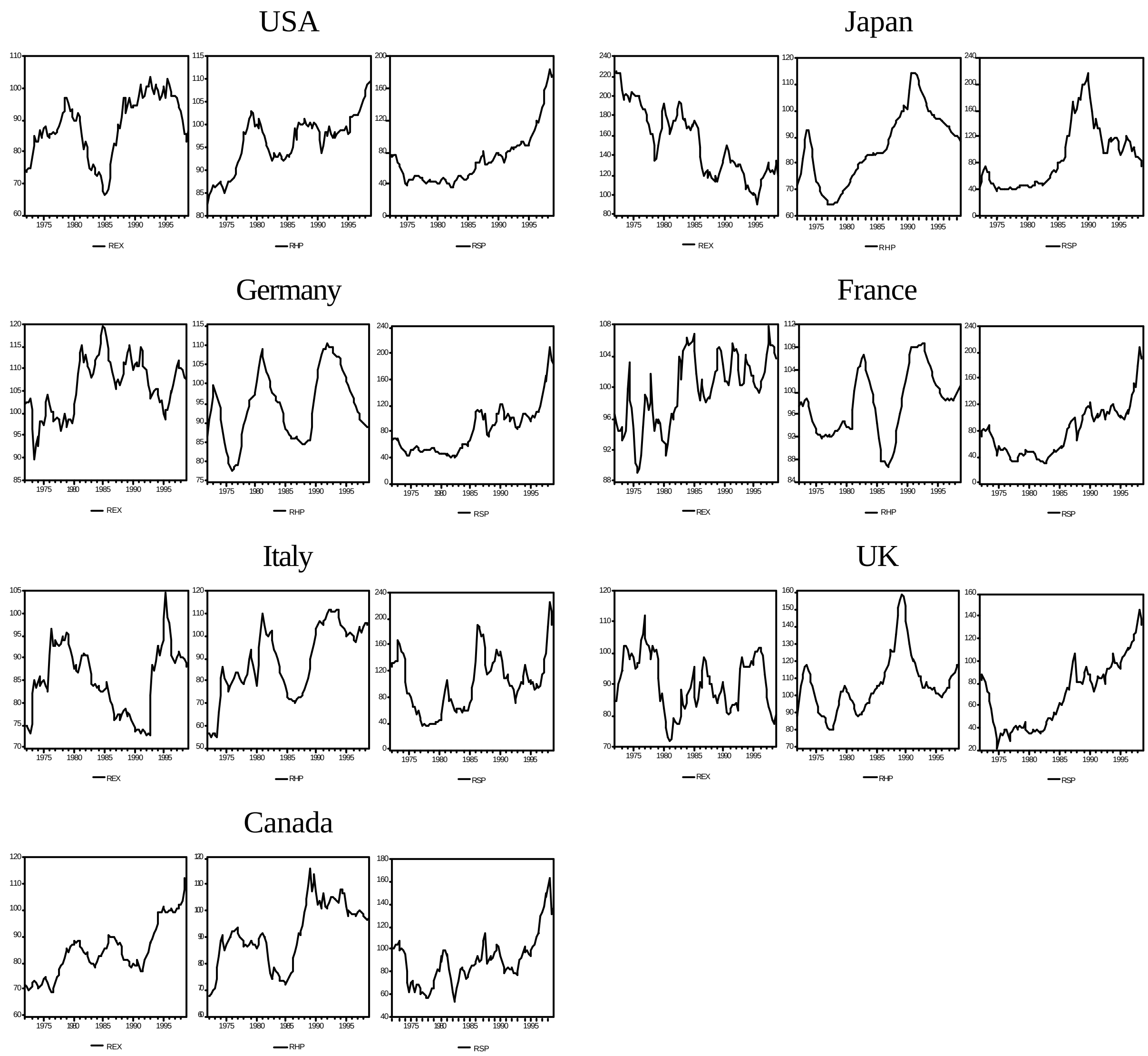
Large-scale macro-econometric models are used by national central banks and governmental organisations, some of which are available for public access. These models are clearly superior to reduced form demand equations and VARs, since they are able to take the structural features of the economy and the interaction of all variables into account. However, large macro-econometric models with an explicit role for house prices are not available for many G7 countries, mainly due to a lack of data availability, a problem that we bypass by resorting to interpolated data in some cases. Since the role of property prices is a centrepiece of this paper, we could not use large-scale macro-econometric models for this reason and thus had to go for the other two options.

The reduced form model consists of a Phillips curve relating CPI-inflation to the output gap and an IS equation relating the output gap to deviations of the short-term real interest rate, the effective real exchange rate, real house prices and real share prices from their long-run trend. The FCI is constructed based on the estimated coefficients in the aggregate demand equation. Alternative FCI weights are derived based on impulse responses in a VAR consisting of the same variables that appear in the reduced-form model.

Modelling long-run trends in asset prices

How should one model long-run trends in asset prices? The short-term real rate of interest rate poses the smallest difficulty in this respect, since it is usually considered to be mean-reverting, with its long-run mean being equal to the long-run growth rate of output. But what about real exchange rates, real house prices and real share prices? A look at Figure 1, which displays the development of these asset prices over 1972-1998 makes clear that the assumption of mean reversion would not appropriately characterise the process that drives these variables. There is a large literature arguing that asset prices are random walks and that it is not possible to identify equilibrium values of asset prices. On the other hand, there is also a large, and growing, literature on how to model equilibrium asset prices. Our view is that it is generally possible to identify periods of misalignment in asset prices if only ex-post, so that we venture to follow the second approach here by assuming that asset prices follow (possibly time varying) determinist but not stochastic trends.

Figure 1: Asset price movements in the G7 countries 1972 - 1998



Note: rex is the effective real exchange rate (units of home goods per unit of foreign goods), rhp is the real house price index and rsp is the real share price index. Data sources are described in the data appendix.

The most commonly used equilibrium concept for real exchange rates, Purchasing Power Parity, implies in its traditional form that real exchange rates are mean reverting[3]. However,

[3] There are several competing concepts to calculate equilibrium exchange rates, most of which require large macro-econometric models. Following all of them up is therefore beyond the scope of this paper. MacDonald (2000) provides a comprehensive survey of all the concepts.

there are also real determinants of the real exchange rates, which could give rise to long-run deterministic trends in real exchange rates. For example, the Balassa-Samuelson effect implies that if a country has higher long-run productivity growth in its tradable goods sector than its trading partners, then its real exchange rate will appreciate in the long-run. Thus, there can be a long-run trend in real exchange rates. For this reason, we model the long-run trend of real exchange rates by regressing the real exchange rate on a constant and a linear trend. We take the same approach for real house prices, based on a related argument. Given that house prices are tied to construction costs, real house prices will follow a long-run deterministic trend since productivity in the construction sector is growing at a slower pace than overall productivity. Thus, the long-run trend in house prices is also modelled by regressing real house prices on a constant and a linear trend. For real share prices the case is more difficult. According to standard asset pricing models, today's share prices should reflect the discounted sum of future real dividends. If the discount rate and the expected growth rate of future dividends is constant, real share prices are given by $P = D/(r-g)$, where D are real dividends, r is the discount rate and g the long-run growth rate of dividends. Since dividends are related to real activity, it would seem to be possible to model long-run trends in share prices also by a linear trend. Visual inspection of real share price movements shows, however, that we are unable to keep track of share prices with a linear trend. Real share prices rather seem to have followed three different regimes over the period 1970-1998. A high, constant level before the first oil price shock in 1974, a low constant level after the first oil price shock and an upward trend since the early eighties. The problem is that the expected long-run growth rate of dividends cannot be assumed to be constant. Adverse supply side conditions, caused by the increase in oil prices and union wage pressure in some countries are likely to have had an adverse effect on expected future dividend growth and thus to have caused the drop in share prices in the early 1970s. Share prices remained at their low levels until the early eighties, when oil prices started to recede and conservative governments in many countries put in place policy measures to improve the supply side of their economies, thus causing an increase in expected future dividend growth. Thus, the long-run trend of real share prices has been time-variable due to the effect of expectations. For this reason we calculate the trend in real share prices using a Hodrick-Prescott Filter with a high smoothing parameter of 10,000, in order to obtain a smooth, time variable trend.

Table 3 reports unit-root test statistics for the derived asset prices gaps. In most cases the null of non-stationarity is clearly rejected, so that valid test-statistics can be obtained from regressions with these gap measures.

Table 3: Unit root tests for asset price gaps

	Interest rate	Exchange rate	House prices	Share prices
USA	-1.88* -3.62***	-2.49** -1.58	-2.09** -1.75*	-4.80*** -3.93***
Japan	-3.01*** -4.85***	-2.92*** -2.20**	-4.23*** -2.99***	-3.11*** -2.99***
Germany	-3.28*** -5.02***	-2.59*** -2.29**	-3.23*** -1.75*	-4.02*** -3.09***
France	-1.91* -3.35***	-3.45*** -3.36***	-3.78*** -1.95**	-4.04*** -3.23***
Italy	-1.87* -3.00***	-2.11** -2.08**	-2.35** -2.54**	-3.59*** -2.37**
UK	-2.30** -4.13***	-2.71*** -2.32**	-3.62*** -2.01**	-4.16*** -3.97***
Canada	-2.15** -3.56***	-1.68* -1.50	-2.21** -1.91*	-3.78*** -3.45***

*Note: The table displays for each country and asset price first the Augmented Dickey-Fuller and then the Phillips-Perron test statistic based on regressions with four lagged differences and a lag truncation of four respectively and without intercepts and trends. *,** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level respectively. The respective critical values are –1.62, -1.94 and –2.58 (MacKinnon, 1991).*

When estimating FCI weights using reduced form equations and VARs, we inherit several potential caveats from the MCI literature (see Eika, Ericsson and Nymoen, 1996 and Ericsson, Jansen, Kerbeshian and Nymoen, 1998). The main criticisms that also apply to our analysis are

- parameter non-constancy
- model dependence of the derived weights
- non-exogeneity of regressors

Parameter inconstancy is potentially a problem in our analysis. Our sample period covers the post-Bretton-Woods period, which could be seen as a single regime with fixed exchange rates in Continental Europe and mainly flexible exchange rates in the rest of the G7. However, our sample period can still be seen as covering more than one regime. The 1970s were characterised by the oil shocks, rather accommodative monetary policy and, in some countries, labour disputes. In the 1980s inflation receded due to a change in the paradigm of macroeconomic policy in the G7 countries with a strong emphasis on stability and, in most G7 countries, on the improvement of supply side conditions. The 1990s saw German re-unification and the bursting of asset price bubbles in Japan, leaving the country in an ongoing depression. Thus, parameter non-constancy is a potential problem for our analysis, which we try to cover mainly by doing standard breakpoint tests.

Model dependence of estimated effects is a potential caveat that applies to any kind of empirical analysis. We were unable to construct (or use) large-scale macro-econometric models nor could we include every variable that possibly affects aggregate demand and inflation, so that the weights we derive certainly depend on the way we specified our model. We hope, however, that the model specification we chose is adequate for the question at stake.

Non-exogeneity of regressors might be seen as a problem that applies with particular force to our analysis, since property and share prices are often characterised as forward looking variables, so that including them as regressors might introduce a simultaneity bias in the estimating equations. However, simultaneity problems may already arise from including interest rates and exchange rates in the analysis. The Central Bank raises interest rates in anticipation of future output gaps, and exchange rates appreciate in anticipation of output gaps, which could trigger rising interest rates. Thus, including property and share prices in the analysis does not introduce a problem that has not potentially been there before. In the literature on monetary policy transmission and monetary policy rules it is always assumed that the information set of the Central Bank can be characterised by lags of endogenous and exogenous variables, so that no simultaneity problem arises. But why should asset markets have information on future output and inflation that are superior to those of the Central Bank? There is no reason why this should be the case, even if stock market investors are always fully rational.

3. Financial Conditions Indices for the G7 countries

FCI weights are derived based on reduced form estimates of coefficients in a simple aggregate demand equation and impulse responses from an identified VAR. The sample period of the analysis is the first quarter 1973 to the fourth quarter 1998, thus covering the post-Bretton-Woods and the pre-EMU era.

Reduced form estimates

In order to assess the importance of financial variables in the conduct of monetary policy, we estimate an extended version of the standard inflation targeting model proposed by Rudebusch and Svensson (1998)[4]. In this framework the economy is modelled by a backward-looking supply or Phillips Curve and a backward-looking demand or IS curve:

$$(1) \qquad p_t = a_1 + \sum_{i=1}^{n1} b_{1i} p_{t-i} + \sum_{j=1}^{n2} b_{2j} y_{t-j} + \sum_{k=0}^{n3} b_{3k} dpo_{t-k} + e_t$$

$$(2) \qquad y_1 = a_2 + \sum_{i=1}^{m1} g_i y_{t-i} + \sum_{j=1}^{m2} l_{j1} rir_{t-j} + \sum_{k=1}^{m3} l_{3k} rex_{t-k} + \sum_{l=1}^{m4} l_{4l} rhp_{t-l} + \sum_{p=1}^{m5} l_{5p} rsp_{t-p} + h_t$$

p is four-quarter inflation in the consumer price index[5], measured as the four-quarter change in the log CPI, y is the percent gap between industrial production and potential industrial production, where potential industrial production is calculated using a Hodrick-Prescott-Filter with a smoothing parameter of 1600. dpo is the quarter-to-quarter change in the world price of oil. This variable acts as a proxy for supply shocks and helped to eliminate heteroskedasticity. rir is the ex-post real short-term interest rate, measured as the short-term money market rate less quarterly inflation. Rex, rhp and rsp are respectively the percent gap between the real

[4] In Goodhart and Hofmann (2000b) we did a similar exercise for a larger sample of industrialised countries. There we demonstrate that in virtually all countries it is not possible to find a significant effect of interest rates on the output gap if the effects of other financial variables are not controlled for.

[5] We used four-quarter-inflation instead of quarterly inflation because year on year inflation is certainly the much more relevant measure of inflation when it comes to policy decisions and quarterly inflation contains a substantial amount of noise that is filtered out when taking four-quarter differences.

effective exchange rate[6], real house prices, and real share prices and their long-run trend values, measured as described above.

The Phillips and IS curves were estimated separately by OLS over the sample period 1973:1-1998:4. The lag order has been chosen by a general-to-specific modelling strategy, keeping all lags between the first and the last significant lag. Tables 3 and 4 report the sum of the coefficients of the exogenous variables with t-statistics in brackets and some diagnostics: the adjusted coefficient of determination ($\overline{R}^2$), a supremum F-test (Sup-F) for unknown breakpoint (Andrews, 1993), a test for first order autocorrelation (LM1) and a test for autocorrelation up to order five (LM5), White's test for heteroskedasticity (H) and a Jarque-Berra test for normality (N). *,**,*** indicates significance of a coefficient or a test statistic at the 10%, 5% and 1% level respectively.

Table 4 reports the estimates of the G7 Phillips Curves. The output gap is significant at the 1% level in all countries, with an estimated effect on inflation ranging between 0.057 (Germany) and 0.14 (UK). The oil price is also significant in all countries, with a particularly strong effect on inflation in Italy and the UK. The estimates of the lagged dependent variables are not reported due to the space constraint. The sum of the lagged inflation terms was not significantly different from 1 in all countries, implying that the long-run Phillips Curve is vertical. The fit and the diagnostics of the estimated equations are satisfactory, with no indication of severe misspecification. Only for Japan and the UK there is some weak evidence of non-normality in the residuals. However, despite the inclusion of the change in the world price of oil we still had to include dummy variables for a couple of large outliers in periods around the oil price shocks in order to eliminate heteroskedasticity and non-normality in some cases. A list of the dummies included for each country can be found in the appendix.

[6] The exchange rate is measured as units of home currency per unit of foreign currency, so that an increase in the real exchange rate is a real depreciation.

Table 4: OLS estimates of the Phillips Curves, Sample 1973:1-1998:4

Country (Lags)	*Coefficient estimates*		*Diagnostics*		
	gap	*dpo*	$\overline{R}^2$ *Sup-F*	*LM1* *LM5*	*H* *N*
USA (1-9,1-3,0-2)	0.066*** (2.92)	0.006*** (4.46)	0.98 2.48	0.01 1.77	37.56 1.92
Japan (1-9,1,2)	0.085*** (3.41)	0.002* (1.77)	0.98 10.63	1.18 7.42	33.33* 4.07
Germany (1-5,1-4)	0.057*** (3.05)	0.002*** (2.79)	0.95 2.11	1.77 5.12	14.71 0.09
France (1-5,1)	0.064*** (3.46)	0.006*** (6.09)	0.99 7.37	2.39 5.57	14.40 0.58
Italy (1-7,1,0-3)	0.096*** (3.60)	0.014*** (6.55)	0.99 5.07	0.92 8.66	30.27 0.73
UK (1-5,1,0-4)	0.14*** (3.07)	0.014*** (5.44)	0.97 4.45	1.23 8.21	33.97 7.33**
Canada (1-6,1-3,0-1)	0.099*** (4.68)	0.002** (1.97)	0.98 2.90	0.19 2.43	28.07 2.78

Note: The dependent variable is four-quarter inflation. Gap is the output gap, dpo the quarterly change in the world price of oil. The table shows the coefficient estimates with t-statistics in brackets. $\overline{R}^2$ *is the adjusted coefficient of determination, Sup-F a test for unknown breakpoint based on Andrews (1993), LM1 and LM5 are test for autocorrelation of order one and up to order five respectively, H is White's test for heteroskedasticity (White, 1980) and N is a Jarque-Berra test for normality. *,** and *** indicates significance of a coefficient or a test statistic at the 10%, 5% and 1% level respectively.*

Table 5 reports the estimates of the IS equations. Except for the real exchange rate in the US, all asset prices enter significantly the IS equations in all countries. Again, due to the space constraint, we do not report the estimates for the lagged dependent variable. The output gap shows significant persistence in all countries, with sum of lagged output gap coefficients being equal to 0.7 on average, but significantly smaller than one. The OECD-output gap enters significantly only in Germany and Italy and was thus eliminated in the other countries. Again, the fit and the diagnostics do not give any indication of misspecification, with only weak evidence for non-normality for Canada and the US. As for the Phillips Curves, well behaved residuals could often only be obtained by including dummy variables for large outliers in the equation, most of which are related to the oil price shocks[7]. A list of the dummies included for each country can also be found in the appendix.

[7] The change in world oil prices itself did not have any significant effect on the output gap.

Table 5: OLS estimates of the demand-equations, Sample 1973:1-1998:4

Country (Lags)	*Coefficient estimates*					*Diagnostics*		
	rir	*rex*	*rhp*	*rsp*	*oecd*	$\bar{R}^2$ *Sup-F*	*LM1* *LM5*	*H* *N*
USA (1-4,1-4,1,1,1)	-0.111* (-1.94)	-0.009 (-0.57)	0.061** (2.13)	0.032*** (2.98)	-	0.92 4.38	0.77 4.15	33.40 5.50*
Japan (1-4,6,1,1-3,1)	-0.109*** (-2.94)	0.022** (2.04)	0.115*** (3.12)	0.016* (1.98)	-	0.93 2.92	2.19 8.17	19.22 2.74
Germany (1-5,1-2,1,1,1)	-0.316*** (-4.02)	0.065*** (2.82)	0.092** (5.09)	0.047*** (4.92)	0.309*** (3.97)	0.84 2.25	0.23 4.69	18.04 0.56
France (1,1-3,1,1,1)	-0.173*** (-4.13)	0.063* (1.71)	0.058*** (2.97)	0.029*** (3.51)	-	0.80 5.76	0.15 2.78	23.17 0.75
Italy (1-4,3,1,1,1-6)	-0.116*** (-3.11)	0.042* (1.75)	0.036** (2.41)	0.03*** (3.52)	0.560*** (5.22)	0.84 4.52	1.70 5.56	37.06 2.34
UK (1-4,1,1,1,1)	-0.066** (-2.39)	0.036*** (2.99)	0.020** (2.13)	0.022** (2.46)	-	0.88 1.29	0.68 9.28	16.35 4.15
Canada (1-4,1,1,1-2,1)	-0.078** (-2.15)	0.041** (2.17)	0.034*** (2.77)	0.033*** (3.22)	-	0.94 2.62	0.04 6.01	23.75 5.27*

Note: The dependent variable is the output gap. Rir is the ex-post short-term real rate of interest, rex is the effective real exchange rate gap, rhp is the real house price gap, rsp is the real share price gap and oecd is the OECD-output gap. The table shows the coefficient estimates with t-statistics in brackets. $\bar{R}^2$ *is the adjusted coefficient of determination, Sup-F a test for unknown breakpoint based on Andrews (1993), LM1 and LM5 are test for autocorrelation of order one and up to order five respectively, H is White's test for heteroskedasticity (White, 1980) and N is a Jarque-Berra test for normality. *,** and *** indicates significance of a coefficient or a test statistic at the 10%, 5% and 1% level respectively.*

VAR estimates

The reduced form model estimated in the previous section is based on a particular view of the transmission mechanism. Asset prices affect the output gap, which in turn affects inflation. However, asset prices may affect inflation also via other channels. Direct effects on inflation may be particularly relevant for the exchange rate and house prices via their effect on the price of imported goods and on the cost of housing respectively. The estimated reduced form equations are in fact nothing but the output gap and the inflation equation from a VAR with exclusion restrictions. In this section we explore an alternative way to estimate FCI weights based on impulse responses of inflation to asset price shocks in an identified VAR. For this purpose we estimate a VAR including the same set of variables that appeared in the Phillips and IS curve: four-quarter inflation, the output gap, the short-term real interest rate, the real exchange rate gap, the real house price gap and the real share price gap. The shocks were identified using a standard Cholesky factorisation, with the ordering (output gap, CPI-inflation, Real house prices, Real exchange rate, Real interest rate, Real share prices) for the Continental European countries. For all the other countries we reversed the ordering of the real exchange rate and the real interest rate. The ordering of the first two variables is fairly standard in the monetary transmission literature: The output gap can affect inflation contemporaneously and both variables do not react immediately to shocks of any of the other variables. House prices are ranked third since house prices are also rather sticky. Share prices are ranked last, since share prices are flexible and can thus be assumed to react immediately to all other variables. The difficult point is the ordering of the interest rate and the exchange rate, since there is a potential problem of simultaneity between the two variables[8]. We ordered the exchange rate before the interest rate for the Continental European countries, because these countries were usually engaged in some fixed exchange rate arrangement, so that it seems to be more appropriate to assume that the exchange rate enters the monetary policy reaction function contemporaneously but reacts with a lag to interest rate shocks. Since this argument does not hold for all the other countries, the exchange rate was ordered after the interest rate there, assuming that interest rates do not respond contemporaneously to exchange rates in those countries. Changing the ordering of interest rates and exchange rates did, however, not significantly affect the impulse responses to the various shocks.

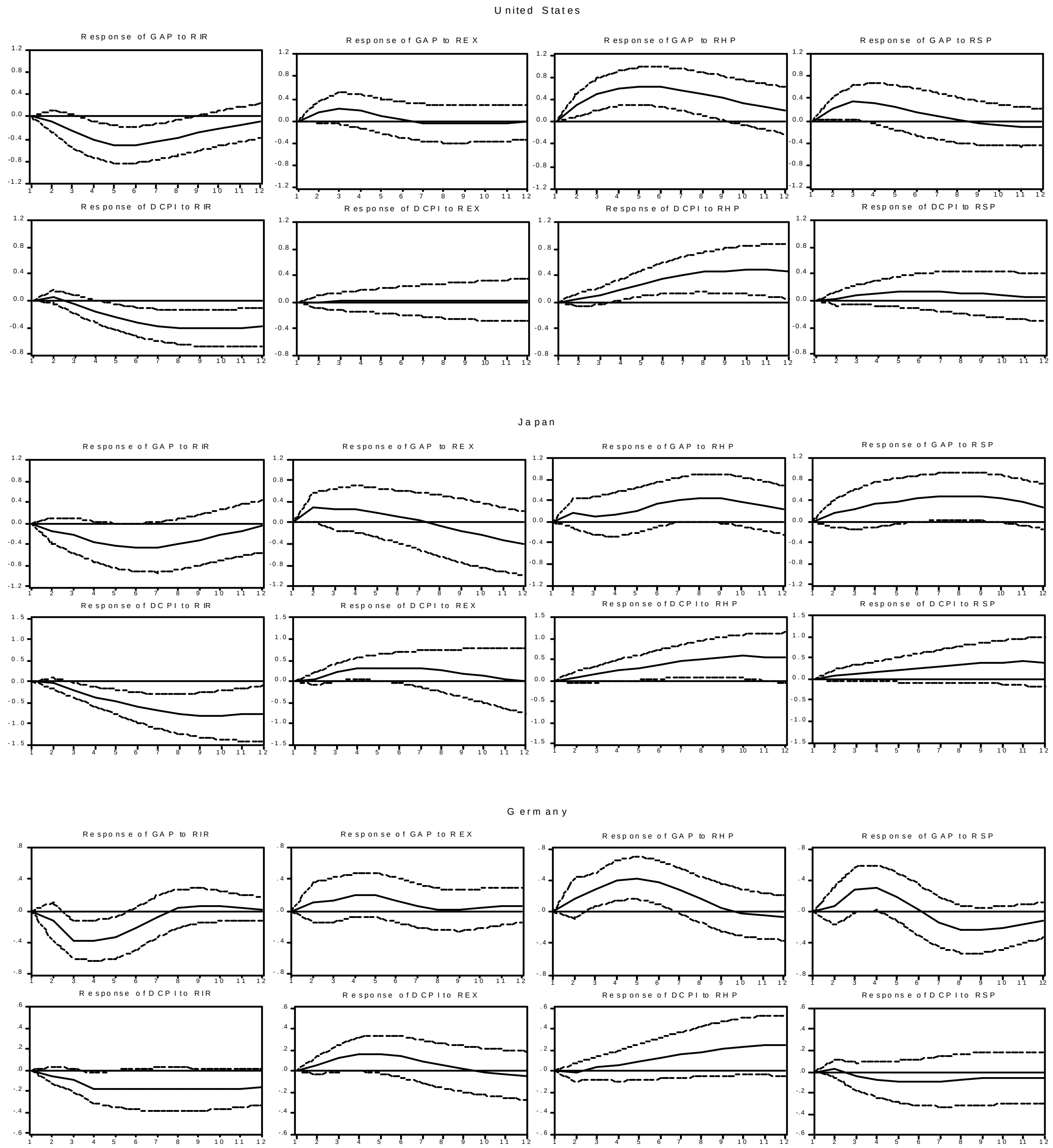
The lag order of the VAR was selected by adopting a general-to-specific strategy, allowing for a maximum lag order of five. A lag order of three was chosen for Canada, Germany, Japan and the UK, for the US two, for Italy four and for France five. The current and the one period lagged change in the world price of oil were included as exogenous variables in the VAR. We also included the set of dummy variables that were included in the reduced form equations and the lagged OECD-output gap in the VARs for Germany and Italy, given its high significance in the reduced form output-gap equations in these countries[9]. Figure 2 shows the impulse responses of the output gap and the inflation rate to a one standard deviation shock to each asset price in a two standard error band. Gap is the output gap, dcpi is CPI-inflation, rhp

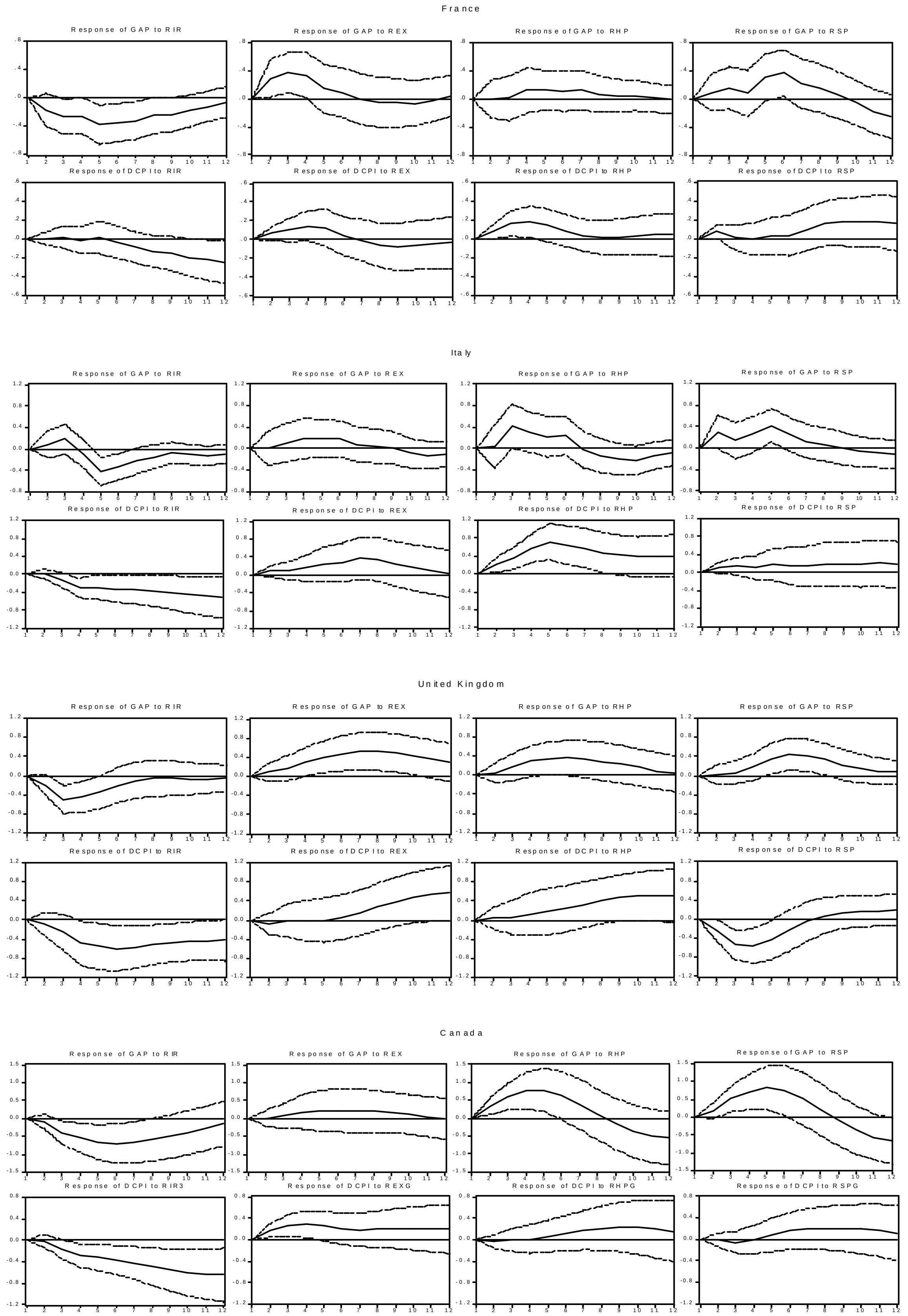
[8] The problem of potential simultaneity between exchange rates and interest rates has not yet been resolved convincingly in the empirical VAR literature, nor is it clear whether the problem is empirically relevant, see Bagliano et al. (1999).

[9]Including the impulse dummies from the reduced form equations proved to be necessary to obtain significant impulse responses for the output gap and the inflation rate. In the Italian VAR the inclusion of the OECD-output gap significantly improved the impulse responses.

are real house prices, rex is the real exchange rate, rir the real short-term interest rate and rsp are real share prices.

Figur2: Impulse responses of the output gap and CPI inflation to asset price shocks



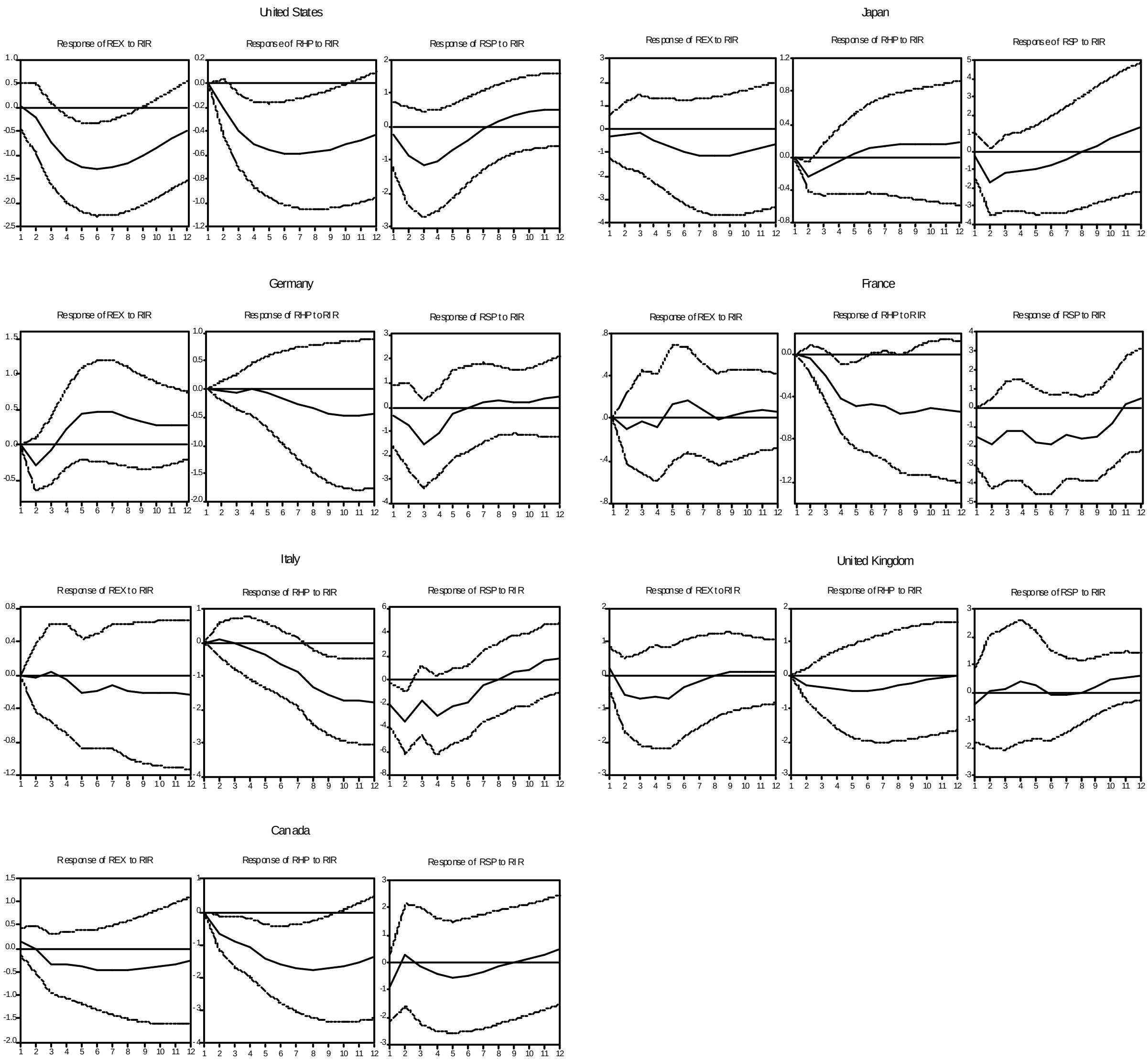


Note: The figure shows impulse responses to one standard deviation shocks in a two standard errors band

The impulse responses reveal that the effect of interest rate shocks on the output gap and CPI inflation is always significant. The responses seem to confirm the theoretical prior that monetary policy shocks first affect the output gap, where the maximum impact is reached after five quarters on average, and then inflation, where the maximum impact is reached after twelve quarters on average. Thus, our results are strongly in line with prior theoretical beliefs about the transmission of monetary policy. The results for the other asset prices are mixed. The impulse responses for the exchange rate are always correctly signed, but we obtain significant output gap responses only for Japan, France and the UK and significant inflation responses for Japan, Germany, the UK and Canada. With the exception of France, the effect of exchange rate shocks on inflation seems to be more pronounced than the effect on the output gap, what can be attributed to the additional direct effect on the CPI via import prices. For shocks to real house prices we also obtain correctly signed impulse responses throughout, which are also significant except for the output gap response in France and the CPI inflation response in Germany and Canada. Except for Germany and Canada the effect on inflation seems to be more pronounced than the effect on the output gap. So, house prices also seem to have a significant direct effect on prices in addition to the indirect effect via the output gap. That result does not come as a surprise, since house prices influence the price of housing, which directly enters the CPI. The results for real share prices are a bit puzzling. We obtain significant positive impulse responses for the output gap in all countries. However, the response of CPI inflation is significantly positive only in Italy. In Germany the response is negative but insignificant, and in the UK the response is even significantly negative. These results are hard to interpret. The only possible explanation we could think of is that there is in fact a forward looking component driving share price shocks. If stock markets expect lower inflation in the future, they also expect lower future interest rates. This could outweigh the positive effect of higher share prices on consumer prices via the output gap.

Another interesting issue is the response of asset prices to interest rate shocks, which are displayed in Figure 3. The impulse responses of asset prices are rather mixed. We obtain no significant impulse responses for real exchange rates, and only three for share prices (Japan, France and Italy). Only for house prices are the results reasonably satisfactory, with five significant (negative) impulse responses.

Figure 3: Impulse responses of asset prices to interest rate shocks



Note: The figure displays impulse responses to a one standard deviation shock to the ex-post short-term real interest rate in a two standard errors band. Rex is the effective real exchange rate, rhp is the real house price index, rsp is the real share price index.

Deriving FCI weights

Based on the estimated coefficients in the IS equation and the VAR impulse responses we can now derive weights for the Financial Conditions Index. For the FCI based on the reduced form estimates we calculate the weight of asset price x by dividing the sum of the absolute value of the estimated coefficients of asset price x by the sum of the absolute value of the estimated coefficients of all four asset prices. The weights for the VAR based FCI are calculated based on the average impact of a one-unit shock to each asset price on inflation over the following twelve quarters. The resulting weights are displayed in Table 6.

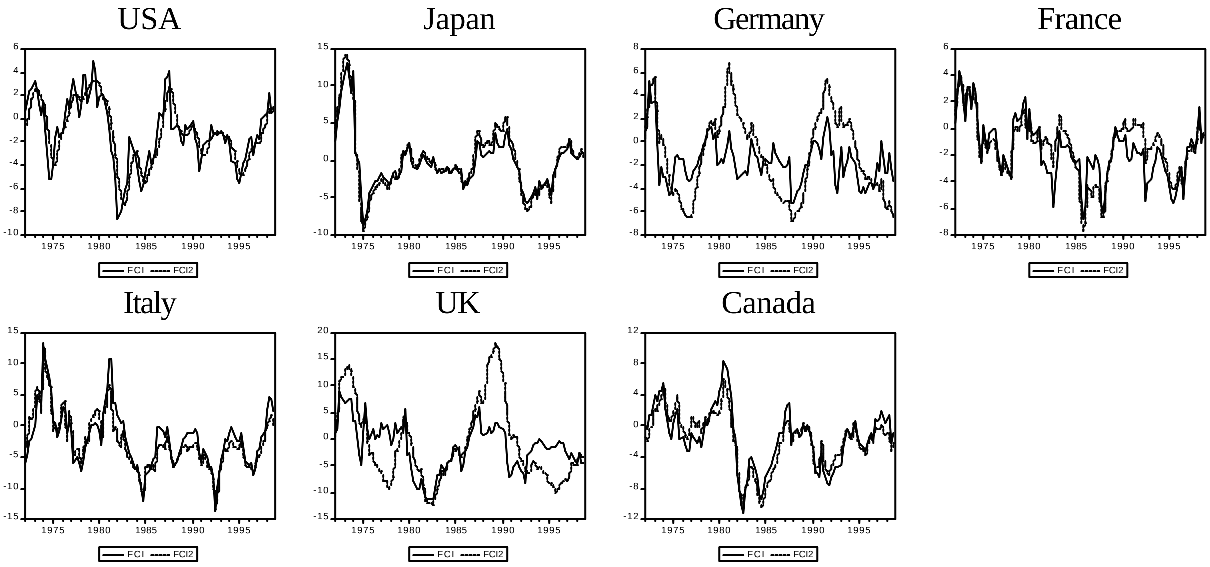
Table 6: FCI weights

	Interest rate	Exchange rate	House prices	Share prices
Canada	0.42 0.59	0.22 0.28	0.18 0.10	0.18 0.04
France	0.54 0.41	0.19 0.04	0.18 0.51	0.09 0.04
Germany	0.61 0.43	0.12 0.08	0.18 0.46	0.09 0.03
Italy	0.52 0.54	0.19 0.13	0.16 0.29	0.13 0.03
Japan	0.42 0.43	0.08 0.04	0.44 0.48	0.06 0.05
UK	0.46 0.45	0.25 0.17	0.14 0.35	0.15 0.03
USA	0.54 0.37	0 0.02	0.30 0.58	0.16 0.03
Average	0.50 0.46	0.15 0.10	0.23 0.40	0.12 0.04

Note: The table shows in each cell first the weight derived from the reduced form estimates and then the weight derived from the VAR impulse responses.

Compared with the weights obtained from the reduced form estimates, we find that house prices get a higher weight and consequently all other asset prices a lower weight in the VAR based FCI. The most straightforward explanation for the higher weight of house prices in VAR based FCI would seem to be the additional direct effect on the CPI via the price of housing. We have seen that in most countries the effect of house price shocks on CPI inflation is more pronounced than the effect on the output gap. However, when looking at the impulse responses it becomes clear that most of the higher weight of house prices in the VAR based FCI derives from the relative stronger effect of house prices on the output gap in the VAR. Figure 4 displays the derived FCIs for each country. Except maybe for Germany and the UK, the reduced form based and the VAR based FCI are very similar.

Figure 4: Financial Conditions Indices for the G7



4. Financial Conditions and future inflation

In-sample evidence

Having derived Financial Conditions Indices, we have to check whether they are of any use to predict future inflationary pressures. As a first simple exercise we look at dynamic correlations of the FCIs with future inflation. Table 7 shows for each country the maximum of the dynamic correlations of the FCIs with future inflation with the respective quarter in brackets. A graph of the full set of dynamic correlations over a twelve quarter horizon is shown in the appendix. The correlation with future inflation is generally quite high, with a maximum G7-average correlation coefficient of 0.55 (0.59) with a lead of seven (three) quarters over inflation. The correlation of the VAR based FCI with future inflation is thus somewhat higher but peaks at an earlier quarter than the FCI base on the reduced form estimates. Thus, the FCIs seem to be useful indicators for future inflation at horizons that are most relevant for policymakers.

Table 7: Maximum correlation of FCIs with future inflation

USA	Japan	Germany	France	Italy	UK	Canada	Average
0.68 (5)	0.60 (5)	0.46 (7)	0.61 (7)	0.57 (3)	0.61 (9)	0.60 (7)	0.56 (7)
0.77 (5)	0.59 (5)	0.65 (1)	0.39 (3)	0.71 (2)	0.59 (5)	0.63 (9)	0.59 (3)

Note: The table shows the maximum dynamic correlation of the FCI with inflation leads, with the respective lead displayed in brackets.

As an additional in-sample exercise we estimate a bivariate VAR with CPI-inflation and each of the two FCIs and do Granger causality tests and compute impulse responses of inflation to FCI shocks.

When carrying out a Granger causality test we examine whether lagged values of the FCI help to predict current CPI inflation, controlling for the information already contained in lagged inflation terms. Table 8 reports the results. The null of no causality is rejected at the 1% level in all cases, so that we conclude that the lagged FCIs contain significant information for future inflation over and above the information already contained in lags of inflation itself.

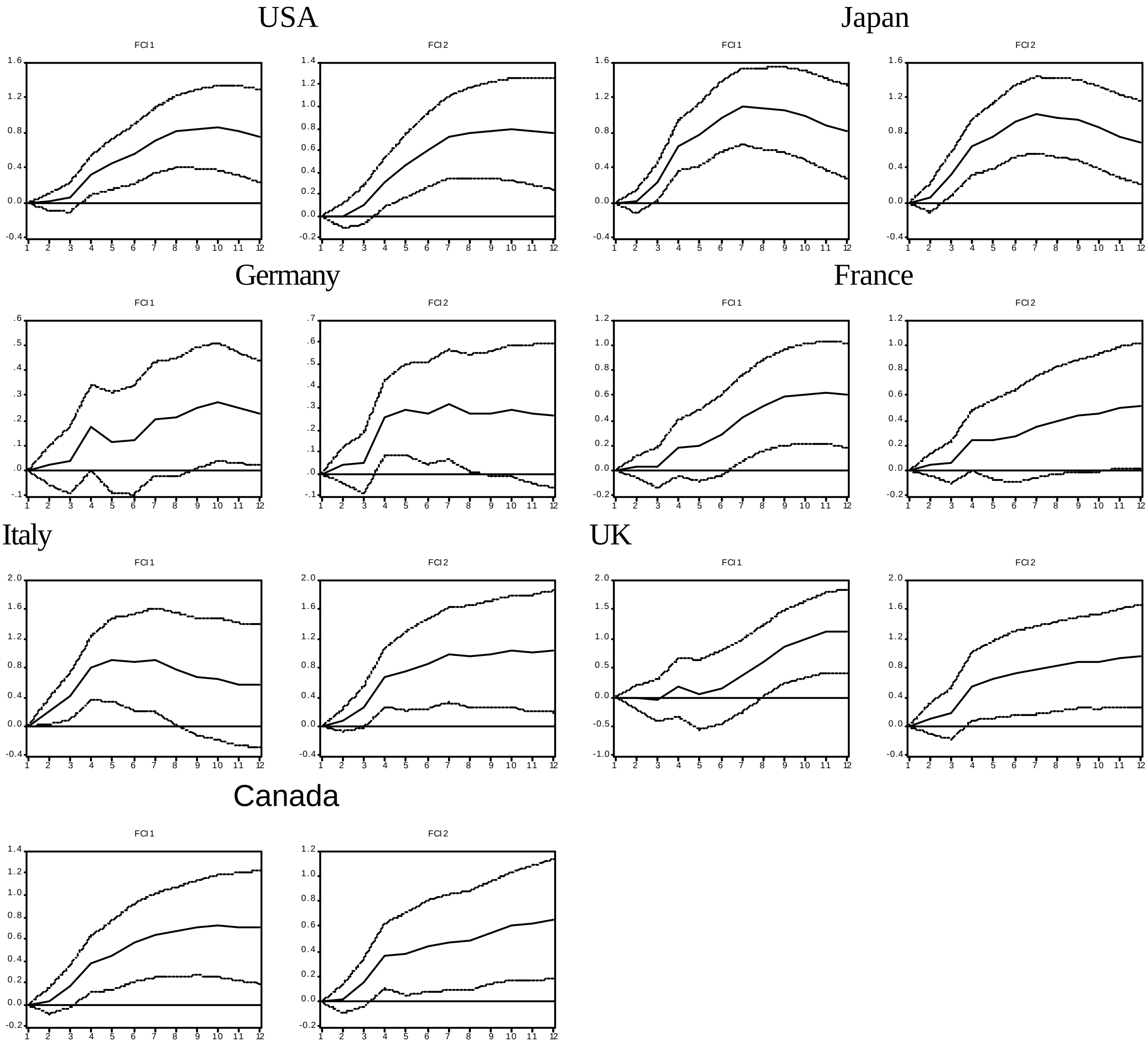
Table 8: Granger causality test for the FCIs

USA	Japan	Germany	France	Italy	UK	Canada
5.71	9.88	6.77	6.49	4.91	6.58	4.19
4.47	5.96	7.03	5.11	8.84	4.18	4.79

Note: The table shows for each country first the Granger causality test statistic for the reduced form based FCI and then the test statistic for the VAR based FCI. Test statistics are based on a regression including five lagged inflation and FCI terms. The 1% critical value is 3.20.

Figure 4 displays the impulse responses of CPI inflation to FCI shocks based on bivariate VARs with a uniform lag order of five. The FCI shock was identified based on a standard Cholesky factorisation ordering CPI inflation first. For both FCIs we find strong and highly significant shock impulse responses of inflation.

Figure 4: Impulse responses of inflation to FCI shocks



Note: The charts display the response of CPI-inflation to a one standard deviation FCI shock in a two standard errors confidence band.

Out-of-sample evidence

Thus, the in-sample evidence implies that FCIs could be quite useful as indicators for future inflation. But how do they perform out-of-sample? It is a common finding in the literature on inflation indicators that good in-sample fit does not guarantee a good performance in out-of-sample forecasting (see e.g. Cecchetti, 1995). We calculate out-of-sample forecasts for the four-quarter inflation rate two years ahead and compare the forecasting performance of the FCIs with that of an purely autoregressive inflation forecast and a random-walk or no-change forecast[10]. Following Stock and Watson (1999, 2000) we choose the h-step ahead projection approach to forecast inflation 8 quarters ahead for the time period 1984:1 – 1998:4. We consider two approaches to model the autoregressive component of the forecasting equation, an autoregressive model and a random-walk model:

(3) $$\Delta cpi_{t+8} = a + g\Delta cpi_t + b(L)FCI_t + e_{t+8}$$ *Autoregressive model*

(4) $$\Delta cpi_{t+8} = \Delta cpi_t + b(L)FCI_t + e_{t+8}$$ *Random-Walk model*

Δcpi is four-quarter inflation, FCI is the Financial Conditions Index and *ß(L)* is a lag polynomial. Out-of-sample forecasts for the period 1983:1 – 1998:4 were calculated by recursively re-estimating the model every period over the sample period starting in 1975:1. The order of the lag polynomial was also updated every period based on the Schwarz-Bayes information criterion. Based on the 64 forecasts we calculated root-mean square errors (RMSEs) to assess the forecasting performance of the FCIs, shown in Table 9. The forecasts obtained from adding a FCI to each of the two models are contrasted with the forecasts obtained from the purely autoregressive model and the no-change forecast. Root-mean square error statistics that are lower than the RMSEs from the benchmark models are in bold.

Table 9:

Root-Mean-Square errors for inflation forecasting two years ahead 1983:1-1998:4

	USA	Japan	Germany	France	Italy	UK	Canada
AR	3.63	2.092	2.137	3.451	5.092	6.131	3.882
AR+FCI1	**2.683**	**2.022**	**2.135**	**3.101**	5.446	**4.64**	**3.328**
AR+FCI2	**2.853**	**1.959**	2.458	3.968	5.172	**5.93**	**3.298**
RW	2.176	1.681	2.022	2.222	2.891	3.232	2.659
RW+FCI1	**2.113**	**1.639**	2.236	**2.184**	3.361	**3.206**	**2.453**
RW+FCI2	**2.038**	**1.62**	2.156	2.375	3.102	3.249	2.876

Note: AR is the autoregressive model for inflation forecasting, RW is the random-walk forecasting model. FCI1 is the Financial Conditions Index obtained from the reduced form estimates, FCI2 is the Financial Conditions Index obtained from the VAR impulse responses.

The table reveals that the Financial Conditions Index based on the reduced form estimates generally performs better out-of-sample than the FCI based on the VAR impulse responses.

[10] This is of course not a real out-of-sample forecasting exercise, since the FCI-weights and the gap measures for output and asset prices were obtained from in-sample analysis. Moreover, in four out of seven countries we are using interpolated data for house prices. For these reasons the out-of-sample forecasting results should be taken with caution.

Except for Italy, the reduced form FCI always outperforms the autoregressive forecast, in most cases, however, only at very small margins. The VAR based FCI outperforms the autoregressive forecast only in four cases. But, all autoregressive forecasts are clearly dominated by the random-walk or no-change forecast. Thus, simply taking this year's inflation rate as the forecast of inflation two years ahead did better than estimating a forecasting model with lagged inflation rates and FCIs every period. However, if the FCIs are integrated in the random-walk framework, we find that the reduced form FCI can improve the random-walk forecast in five countries, the VAR based FCI only in two countries. Thus, the out-of-sample forecasting performance of the FCIs is rather disappointing, since we do not clearly outperform the simple no-change forecast. This finding confirms the conclusion of previous studies (Stock and Watson, 1999, 2000, Cecchetti, 1995, and Cecchetti, Chu and Steindel, 2000) that inflation is very hard to forecast out-of-sample.

5. Conclusions

An MCI, a weighted average of a short-term interest rate and the exchange rate, as an indicator or even operating target in the past has been used by some Central Banks, and strongly criticised in large parts of the academic literature. In this paper we extend the MCI concept to a concept of Financial Conditions by adding house and share prices to the analysis. We analyse the predictive power of asset prices for future output gaps and CPI inflation in the G7 countries for the post-Bretton Woods period based on a simple, reduced form model of the economy and an identified VAR. Prior to the estimation asset prices were de-trended using standard de-trending techniques in order to obtain stationary regressors. The results from the reduced form regressions reveal that all asset prices significantly affect the output gap except for the real exchange rate in the US. From the identified VAR we obtain significant impulse on the output gap and CPI inflation in all cases for the interest shock and in the majority of the cases for the house price shocks. For the exchange rate impulse responses are always correctly signed, but significant only in about half of the cases. The responses to equity price shocks are rather puzzling, being always significant for the output gap, but generally insignificant or even wrongly signed for CPI inflation. This puzzling result might possibly be due to a forward looking element in share price movements.

From the estimated coefficients of the reduced form model and the impulse responses of inflation to the asset price shocks in the VAR we derive two Financial Conditions Indices for each country. The main difference between the two indices is the significantly higher weight for house prices in the VAR based FCI. It appears that both FCIs are fairly useful to predict future inflation in-sample, a finding that is, however, not fully confirmed out-of-sample.

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Appendix

The share of housing assets in household total assets

	1970	1980	1990	1995	1998
United States	22	27	27	23	21
Japan	10	14	8	10	10
Germany	-	-	34	34	32
France	34	44	43	42	40
Italy	36	40	37	35	31
United Kingdom	-	40	44	33	34
Canada	21	22	23	22	21

Source: OECD Economic Outlook 68, December 2000, Table VI.1

The share of equity in household net financial wealth

	1970[1]	1980[2]	1990	1994	1997
United States	-	13.8	12.1	16.0	20.7
Japan	24.7	19.0	26.1	16.1	-
Germany	12.1	5.3	6.1	6.0	9.0
France	-	14.1	22.2	14.6	13.4
Italy	1.9	2.5	7.8	7.3	13.0
United Kingdom	23.0	17.9	23.0	23.1	26.1
Canada	39.9	38.5	34.2	38.4	42.6

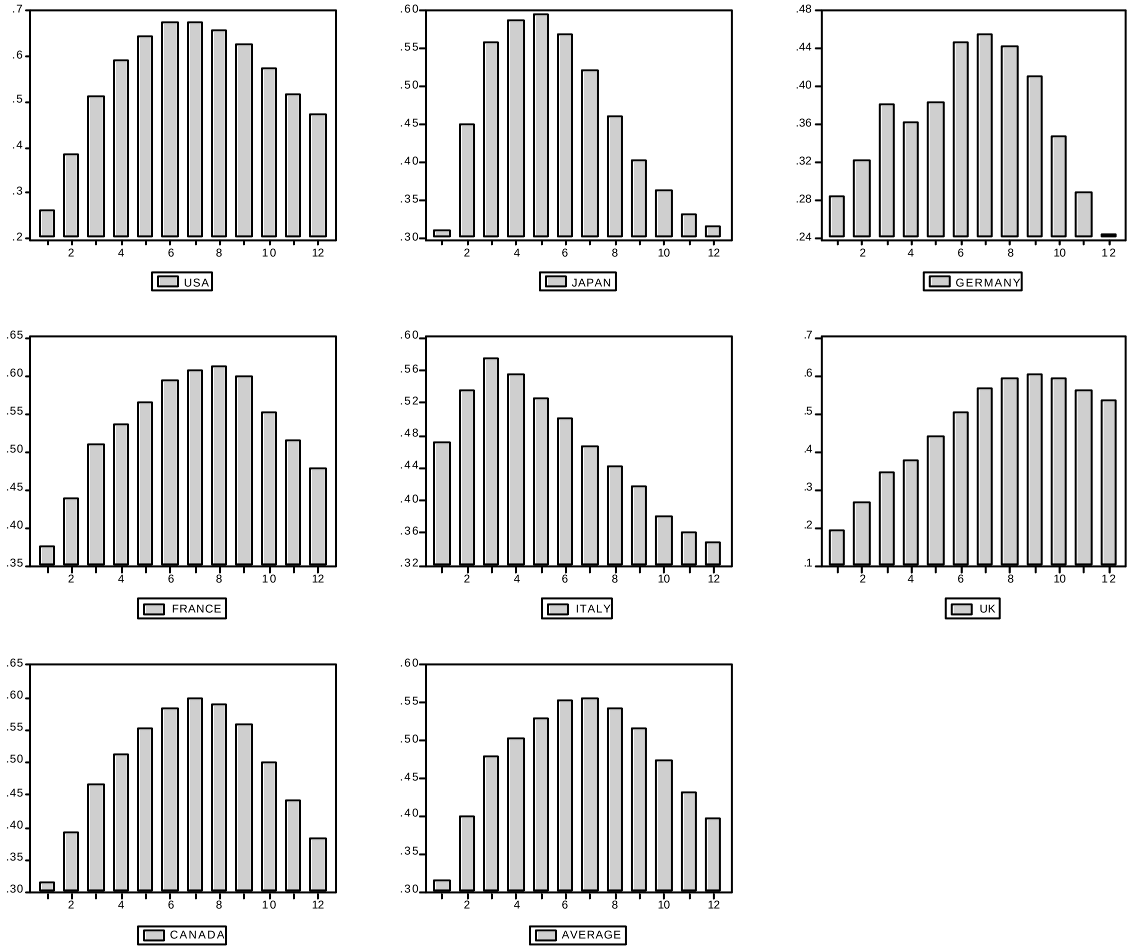
1. 1975 for Italy and the United Kingdom.
2. 1985 for France.

Source: Annex Table 58 in *OECD Economic Outlook No. 64; Boone et al. (1998)*.

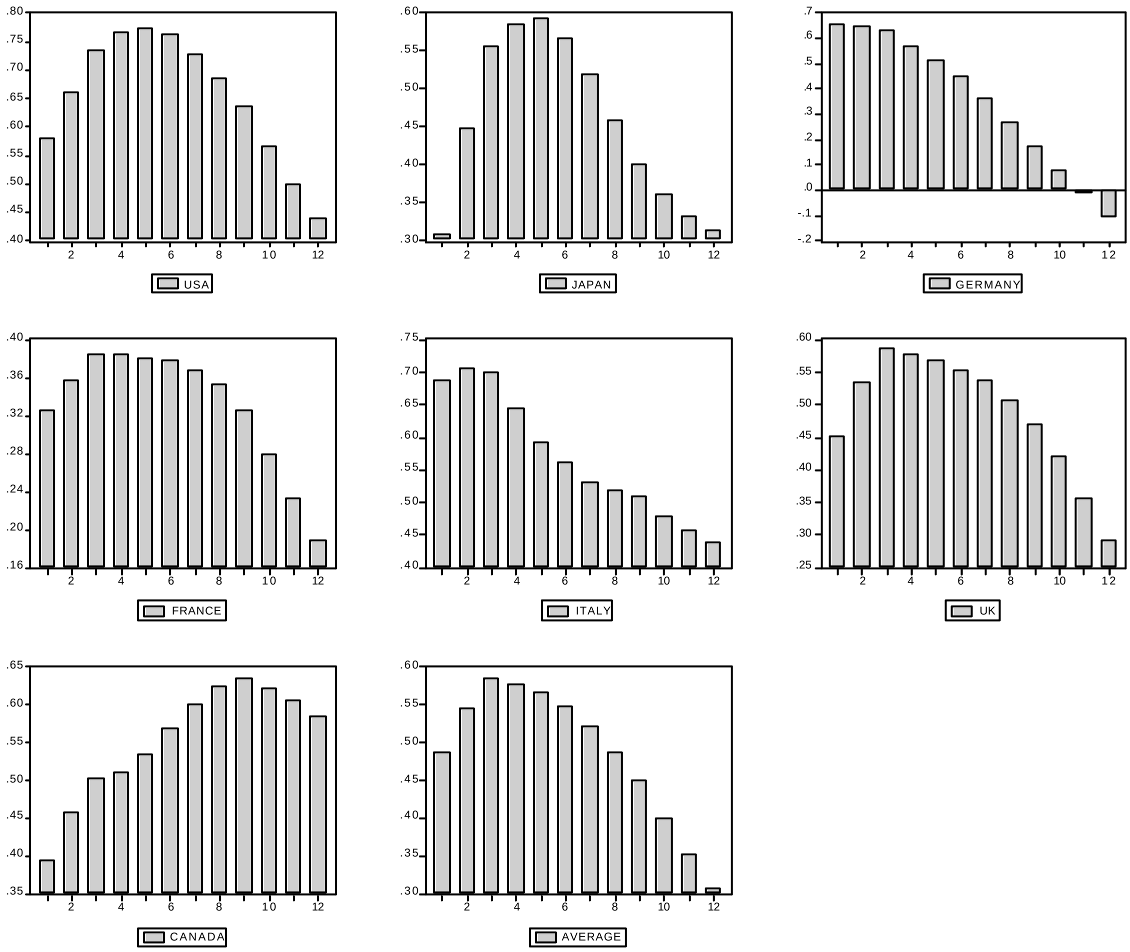
Dummies included in the inflation and the output gap equation

	Inflation equation	**Output gap equation**
USA	-	75:1, 80:2
Japan	74:1	74:3, 74:4, 75:1, 93:4
Germany	74:1, 93:1	-
France	80:1, 82:3	74:4, 75:23
Italy	76: 1/2	74:4, 75:1, 79:1/2, 81:3
UK	75:2, 75:3	74:4, 75:2, 75:3, 79:2, 82:2
Canada	-	75:1, 80:2, 83:1, 85:1

Dynamic Correlations of the reduced form based FCI with future inflation



Dynamic Correlations of the VAR based FCI with future inflation



Data Appendix

Consumer Prices

Consumer Price Index, IMF International Financial Statistics, series code 64; OECD Main Economic Indicators for Germany.

Industrial Production

Industrial Production Index, IMF International Financial Statistics, series code 66..c.

World price index of petroleum

IMF International Financial Statistics

Real Exchange Rates

Effective real exchange rate, OECD Main Economic Indicators.

Short-term interest rates

Overnight money market rates for France, Germany, Italy, Japan and the USA, IMF International Financial Statistics, series code 60b; three months commercial paper rate for Canada (IMF, series code 60bc) and three months interbank rate for the UK (BIS).

Share prices

Share price index, IMF International Financial Statistics, series code 62, OECD Main Economic Indicators for the USA.

Property Prices

Residential property price index from national sources as shown in Appendix-Table 1. For France and Germany we had only annual data, for Italy and Japan we had semi-annual data. These were converted to quarterly frequency by linear interpolation assuming an ARIMA(0,1,0).

Sources of residential property prices series

	Series	Source
United States	Median sales prices of existing one-family houses	National Association of Realtors
Japan	Nation-wide land price index	Japan Real Estate Institute
Germany	Average sales price of dwellings and terraced houses in Frankfurt, Munich, Hamburg and Berlin	Ring Deutscher Makler
France	BIS estimate of residential property prices for the whole country	INSEE and Chambre des Notaires
Italy	House price index for the whole country	Bank of Italy/ 'Il consulente immobiliare'
United Kingdom	Nationwide Anglia House Price Index	Datastream
Canada	Multiple listing service price index of existing homes	Central Bank