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# **Triggers for countercyclical capital buffers**

**Karlo Kauko**

*The opinions expressed in this paper are those of the authors and do not necessarily reflect the views of the Bank of Finland.*



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Mika Pösö

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Postal address  
PO Box 160  
FI-00101 HELSINKI, FINLAND

Street address Snellmaninaukio  
Phone +358 10 8311  
Fax +358 9 174 872

Email  
firstname.surname@bof.fi  
www.bof.fi

Swift SPFB FI HH  
Business ID 0202248-1  
Domicile Helsinki

# 1 Introduction

Banking is said to be inherently procyclical; during good times banks loosen their credit standards, fuelling the boom. During recessions banks become reluctant to grant loans, which slows down economic activity and may turn a recession into a depression. Literature on these issues is reviewed by Gordy and Howells (2006) and Drumond (2009). Various proposals have been made to reduce the procyclicality of banking. The discussion has been vivid especially in the aftermath of the global banking crisis (see e.g. Hukkinen and Kauko 2011).

The so-called countercyclical capital buffer system will be introduced in different parts of the world, for instance in the EU as a part of CRD IV. When the credit market seems overheated, regulators could impose additional capital requirements on banks<sup>1</sup>. This additional requirement would be country specific and depend on the home country of the borrowers, not on the location of bank headquarters. The buffer would protect banks against excessive accumulation of risks and make them more resilient during recessions.

The decision to require the additional capital buffer would depend on regulators' discretion, but some objective criteria would be used as a guideline. For instance, the article 126 of the EU draft directive (July 2011) on capital adequacy requirements refers to "the deviation of the ratio of credit-to-GDP from its long-term trend". This wording is probably based on analyses by Drehmann et al (2010) and the Basel Committee guidance published in December 2010. These authors tested some potentially interesting variables, namely the trend deviation of the credit to GDP ratio, credit growth, GDP growth, property prices and a few variables related to bank profitability. The authors proposed that the decision should be based on the trend deviation of the credit-to-GDP ratio; whenever the loan stock relative to the nominal GDP seems significantly larger than what could be considered normal in the light of recent history, additional capital should be required. Drehmann et al (2011) reached a similar conclusion. The trend deviation should be estimated by applying the Hodrick-Prescott filter to the available data, using an exceptionally high value for the "stickiness parameter" lambda. This proposal has been criticised by Repullo and Saurina (2011), who reached the conclusion that at least in major advanced countries this trigger would typically induce

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<sup>1</sup> See EU Council document 2011/0202 – 2011 /0203; 18 Nov 2011

regulators to impose the capital requirement when GDP growth is low or even negative, because the indicator cannot differentiate between excessive credit growth and recessions.

This paper presents two alternative ways to derive a suitable trigger indicator from data on the loan stock and the GDP. The ability of these two indicators to predict major problems in the banking sector seems to outperform the predictive power of the original proposal by Drehmann et al. in at least the recent international financial crisis.

This paper does not discuss any other aspects of the countercyclical capital buffer as a macroprudential tool. For instance, the efficiency of additional capital requirements as a policy tool and related legal and organisational issues are beyond the scope of this paper.

## 2 Trend deviation of credit to GDP

The indicator proposed by Drehmann et al (2010) is derived from the credit-to-GDP ratio. A central characteristic of the original ratio is its non-stationarity and lack of mean reversion. The relative loan stock has been growing in a trend like yet irregular manner for decades, and there seems to be no clearly observable upper limit to the amount of financial intermediation in the economy. The tendency of financial markets to grow faster than the rest of the economy is sometimes referred to as “financial deepening”, especially in discussions on development economics (See e.g. Apergis & al 2007).

The presence of unit roots in the credit-to-GDP ratio was tested with panel unit root tests with annual 1986-2010 data for three separate country groups, namely major advanced economies, Nordic countries and a rather arbitrary group of developing countries. The data included bank loans only, whereas the proposal by Drehmann et al (2010) was to use a broad group of possible loans, including lending by non-banks.<sup>2</sup> The unit root hypothesis is clearly consistent with the data for each of these country groups separately. (See appendix 1, table 2) This makes the credit to GDP ratio a problematic variable. If the variable can remain at any level for lengthy periods of time, and if the latest observed value tells nothing about likely future changes, the trigger is not particularly useful.

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<sup>2</sup> All the international data in this paper are from World Bank. Bank lending = domestic credit provided by the banking sector; includes all credit to various sectors on a gross basis, with the exception of credit to the central government, which is net. Some of the variables used in different analyses are calculated using World Bank data as inputs.

According to the proposal by Drehmann et al (2010) and the draft EU directive, a proxy for the trend deviation of the credit-to-GDP ratio should be used as the trigger. The idea is probably based on the implicit assumption that at each moment of time, the credit to GDP ratio has got a relatively stable equilibrium value, and the trend deviation is due to cyclical oscillations around the equilibrium. However, the credit-to-GDP variable does not seem to be characterized by any regular tendency to reversion (see appendix 2), at least not in developed countries, which is inconsistent with the hypothesis of a smoothly developing equilibrium and cyclical oscillations around it. Nevertheless, using a proxy for the trend deviation rather than the original data imposes mean reversion on the data, which is a highly desirable property for the trigger variable. Estimating the trend deviation is a completely possible yet somewhat non-standard way to derive a mean-reverting variable from a unit root process.

According to the proposal by Drehmann et al (2010), the trend should be extracted by using the Hodrick-Prescott filter. An essential detail of the method is the value of the “rigidity parameter”  $\lambda$ . With quarterly data, the value 1600 is often used in different contexts, but Drehmann et al (2010) concluded that much higher values, either 125 000 or even 400 000, are preferable. Such high values make the trend almost linear, at least with short samples.

From the point of view of policy decisions, nothing but the last observation matters. Unfortunately this method cannot calculate the final value of the trend value for the last observation. Drehmann et al use the term “one-sided filter” when they estimate the trend value for each observation separately without using data for later moments of time. The inability to take the future into account is no major problem if the initial estimate is reasonably reliable, and does not change much when data on subsequent periods becomes available. However, as the following analyses reveal, the initial estimate is highly provisional, at least if the trend is estimated with a high value of  $\lambda$  and without several decades of data. Paradoxically, this inaccuracy, or the use of the “one-sided” filter, may even be desirable.

At least ex post, it seems that 2006 would have been a good moment to impose the countercyclical buffer requirement in many jurisdictions. A simple test was run with annual data on 25 countries.<sup>3</sup> The Hodrick-Prescott filter was run twice ( $\lambda=488$ , which corresponds to 125 000 in quarterly data<sup>4</sup>), first with annual data from 1990-2006, then with

<sup>3</sup> Argentina, Austria, Belgium, Chile, Spain, Netherlands, Hong Kong, Ireland, Italy, Japan, South Korea, Greece, Poland, France, Sweden, Germany, Singapore, Finland, Denmark, UK, USA, Hungary, New Zealand, Mexico, Portugal.

<sup>4</sup> See Ravn & Ulig 2002.

data from 1990-2010. No weighting is used in calculating the averages. Interestingly, the correlation between the two trend deviation estimates for 2006 is 0.62, clearly positive, but much weaker than what one would expect. Annual observations were used because there is no quarterly data in the database.

- Using the shorter data, one reaches the conclusion that the credit-to-GDP ratio was, on average, 3.2 percentages above the trend value in 2006, consistently with the conventional view that lending booms were more commonplace than credit crunches.
- Using estimates based on the somewhat longer sample (1990-2010), one reaches a different conclusion. The credit-to-GDP ratio was, on average, 2.8 percentages *below* its trend value in this cross-national sample in 2006, as if a credit crunch would have been more commonplace than a credit boom in 2006

The drop in the 2006 trend deviation is due to the weak development of GDP in 2006-2010, which increased the credit-to GDP ratios in many countries during the financial crisis. In the sample of 25 countries, the non-weighted average of loan-to-GDP ratios was 127.8 % in 2006 and 152.2 % in 2010.

If the trend had been estimated with longer series on history, adding a few observations would have had much less impact on estimated 2006 trend values, especially with the very high value of the rigidity parameter  $\lambda$ . The latest observations would have little impact on the almost linear trend and initial estimates would be hardly affected by a few additional years of data. Nevertheless, one may question how meaningful it would be to add data from the early 1980s in the sample, because financial markets were tightly regulated in many countries and the driving forces of the credit stock were quite different from those of liberalised markets.

Finland may be an excellent example of very strange phenomena. Due to the sharp collapse of output in 2009, the nominal GDP in 2009 was almost 7 % lower than in 2008. The loan stock, instead, continued its growth, yet much slower than before, and the H-P filter residual indicated an excessive lending boom in 2009, which is an absurd conclusion. A similar example is the year 1991; due to the collapse of output during the exceptional depression of the early 1990s, the credit-to-GDP ratio indicated that credit growth was excessive.

These problems are closely related to the discussion by Repullo and Saurina (2011); the credit-to-GDP gap is driven by both the GDP and the loan stock. A slow GDP growth rate, let alone a negative one, is interpreted as excessive loan growth by the credit to GDP ratio. In

the sample of Repullo and Saurina, The H-P residual suggested by Drehmann et al (2010) is negatively correlated with GDP growth, which might induce policy makers to impose the capital requirement during recessions.

## 3 Differenced credit to GDP

### 3.1 Two versions of an alternative trigger indicator

An ideal trigger mechanism should satisfy at least the following principles:

1. The indicator should have predictive power; it should make as few type 1 errors (alarm not followed by a crisis) and type 2 errors (crisis without alarm) as possible,
2. a sudden fall in the GDP should not be interpreted as a sign of excessive credit growth,
3. the indicator should be stationary (unless there are major structural changes in the environment affecting financial stability, in which case the change should be reflected in the indicator)
4. the indicator should be resistant to structural changes; changes in, say, banking legislation or business practices, should not have a drastic impact on the level which should be interpreted as an alarm.

An obvious way to weaken the impact of a sudden fall of the GDP on the indicator is to smooth the GDP. Instead of using the latest data, one could use the backwards looking moving average. In the following calculations, the five years moving average is used, even though using a four or six years moving average would probably be equally justified.

Most non-stationary economic variables become stationary by taking the difference, and the credit to GDP ratio is no exception. A few panel unit root tests were carried out, and the stationarity of the differenced ratio seems universal. (See appendix 1, table 3) An advantage of taking differences instead of estimating the trend deviation is that the final value of the variable does not change in the future when new data become available, except if provisional statistics are significantly revised.

By combining the principles 1 and 2, one can present two alternative ways to derive an indicator  $X$  from the data on nominal GDP ( $Y$ ) and the loan stock ( $L$ ).

#### Equation 1. The first indicator

$$X_t = \Delta \left[ \frac{5L_t}{\sum_{i=0}^4 Y_{t-i}} \right] = \left[ \frac{5L_t}{\sum_{i=0}^4 Y_{t-i}} \right] - \left[ \frac{5L_{t-1}}{\sum_{i=1}^5 Y_{t-i}} \right]$$

#### Equation 2. The second indicator

$$X_t = \left[ \frac{5 \Delta L_t}{\sum_{i=0}^4 Y_{t-i}} \right]$$

Subscripts denote years.

Both proxies for the state of the loan market ( $X$ ) have at least one desirable characteristic: at least simple test statistics indicate they are stationary in a broad range of different countries (see appendix 3). Being significantly under or above the long-run average implies that the situation is exceptional and will be short-lived due to an inevitable correction. Whether the correction takes place as a crisis or not is a different issue, but imbalances exist.

During periods of accelerating inflation these indicators may be misleading. The moving average of the nominal GDP may seem low because the output of the past was sold at a lower price level. This may be a problem of the 2<sup>nd</sup> indicator, but less so if the first indicator is used. Hence, inflation could be taken into account when one evaluates which values of the 2<sup>nd</sup> trigger should be regarded alarmingly high.

### 3.2 Performance of alternative indicators in a cross-national sample

Four indicators were calculated for a number of countries. These indicators are two versions of the Hodrick-Prescott residual of the credit-to-GDP based on data for 1990-2006, the first indicator and the second indicator. The Hodrick-Prescott residual is calculated using two possible values of  $\lambda$ . The first value is 488; with annual data it should correspond to the value 125 000 applied on quarterly data, which according to Drehmann et al (2010, p 29) performs well. The second parameter value is 100, which is often used with annual data in different

contexts. These data for 2006 are presented in table 1. The choice between the two possible values of lambda seems to affect the absolute value of estimated credit-to-GDP gaps, but the ranking of countries remains broadly unchanged. In addition to the original value, each indicator is presented in a standardised form. Standardised indicators are derived from original ones by subtracting the mean in 2006 from each observation and by dividing the result by the standard deviation of the variable across countries. Each standardised indicator has got mean zero and standard deviation one, which facilitates comparisons.

Table 1. Different indicators in 2006, bank loans only

	H-P Trend deviation, $\lambda = 488$	H-P Trend deviation $\lambda = 100$	Indicator 1 (in %)	Indicator 2 (in %)	Standardi-sed H-P Trend deviation $\lambda=488$	Standardi-sed H-P Trend deviation $\lambda=100$	Standardi-sed Indicator 1	Standardi-sed Indicator 2
Argentina	-13.56	-11.88	-8.10	-0.33	-1.56	-1.84	-1.68	-1.24
Austria	2.33	2.63	0.75	6.63	-0.08	0.07	-0.57	-0.54
Belgium	-2.61	6.21	5.91	10.65	-0.54	0.54	0.07	-0.13
Chile	-6.76	-5.98	2.81	13.55	-0.93	-1.07	-0.31	0.16
Denmark	5.59	-2.94	14.61	22.35	0.22	-0.67	1.16	1.06
Eire	23.34	18.00	23.82	36.00	1.88	2.09	2.31	2.44
Finland	14.63	7.84	5.52	9.13	1.06	0.76	0.02	-0.28
France	6.13	4.77	7.51	11.99	0.27	0.35	0.27	0.01
Germany	-13.56	-7.66	-3.38	0.22	-1.56	-1.29	-1.09	-1.19
Greece	6.27	3.35	3.00	10.95	0.28	0.16	-0.29	-0.10
Hong Kong	-8.22	-6.34	-6.39	0.76	-1.07	-1.11	-1.46	-1.13
Hungary	17.35	10.80	6.39	11.26	1.32	1.14	0.13	-0.07
Italy	6.36	3.98	4.67	8.55	0.29	0.25	-0.08	-0.34
Japan	-11.88	-8.90	-9.74	-6.65	-1.41	-1.45	-1.88	-1.88
Korea	2.27	1.23	7.40	12.88	-0.09	-0.12	0.26	0.10
Mexico	2.50	2.86	3.61	6.94	-0.07	0.10	-0.21	-0.51
Netherlands	-0.03	-1.74	2.86	10.34	-0.30	-0.51	-0.31	-0.16
Poland	2.72	2.34	4.84	7.78	-0.05	0.03	-0.06	-0.42
Portugal	0.22	-0.14	12.22	17.83	-0.28	-0.30	0.86	0.60
Singapore	-13.40	-8.09	0.31	6.98	-1.55	-1.34	-0.63	-0.50
Spain	24.14	16.35	20.53	32.63	1.95	1.88	1.90	2.10
Sweden	13.71	8.75	6.43	12.34	0.98	0.87	0.14	0.04
United Kingdom	11.99	7.25	10.88	19.68	0.82	0.68	0.69	0.78
United States	4.13	4.03	11.23	24.06	0.08	0.25	0.74	1.23
New Zealand	6.94	6.03			0.35	0.52		
Average	3.22	2.11	4.67	10.72				

Standardised indicator for each country  $i$  calculated in the following way.

$$x_{i,standardised} = \frac{x_i - \bar{x}}{StDev(x)}$$

Where  $x_i$  is the original indicator,  $\bar{x}$  = the mean of the indicator in the 2006 sample and  $StDev(x)$  is the standard deviation of the indicator across countries in 2006.

Casual observations indicate that each indicator would have had some predictive power in forecasting cross-national financial instability. The trend deviation indicators had alarmingly high values in Spain, Ireland and Hungary, which seems reasonable *ex post*. On the other hand, the indicator was only moderate for the U.S. Higher values were observed in e.g. Finland and Sweden, even though these countries suffered much less from the crisis. The two different versions of the differenced relative loan stock would have detected excessive credit growth in Ireland, Spain and, perhaps surprisingly, Denmark. The very rapid credit growth in Denmark had continued for many years, and the H-P filter interprets this as a persistent trend.

The predictive abilities of indicator 1, indicator 2 and the H-P filter residuals were tested in a more systematic way with a relatively limited cross-national sample of 24 countries. The focus is solely on the development of credit quality during the recent financial crisis.

The countries are those listed in footnote 1 and table 1, except that New Zealand had to be excluded because of lack of data. The explained variable was the logarithm of the ratio of NPLs to total loans in 2009<sup>5</sup>. The 2006 value of the explained value was used as a control variable. Explanatory variables are those of table 1.

When controlled for the relative amount of NPLs in 2006, the 2nd indicator has got the best predictive ability. Despite the very limited number of observations, this indicator is extremely significant in this simple cross-sectional OLS analysis. The first indicator is also highly significant, except if controlled for the second indicator. The H-P filter residuals are significant, but they have less predictive power, and they lose their significance if either indicator 1 or indicator 2 is included in the analysis. (See appendix 4)

### 3.3 An application to historic Finnish data

The predictive power of indicators 1 and 2 was tested in section 3.2. and in appendix 4. A central problem of this analysis is obviously the very short time dimension and focus on one particular crisis. A good indicator should be universally valid, be robust to relatively large structural changes and be able to predict different kinds of crises. Whether a given value is alarmingly high or not should depend as little as possible on the era and the country. If mergers among financial institutions, financial innovations and changes in banking legislation change the dynamics of the indicator in an unknown way, the indicator is of little use. Such

<sup>5</sup> Source of data; IMF Global Financial stability report, statistical appendix, p 56-57 and fsi.imf.org

structural changes take place continuously. There would always be a very limited amount of data collected under circumstances relatively similar to those that currently prevail, implying that one cannot test the predictive power of potential trigger variables with a sufficiently large sample.

Suitable annual data on credit and loan aggregates for Finland were readily available<sup>6</sup> for a very long period of time. The indicators 1 and 2 were calculated for years 1905-2010. Both indicators appear stationary (see appendix 5).

Finland experienced three banking crises during this period, one in the early 1920s, a second one in the early 1930s and a third one in the early 1990s.

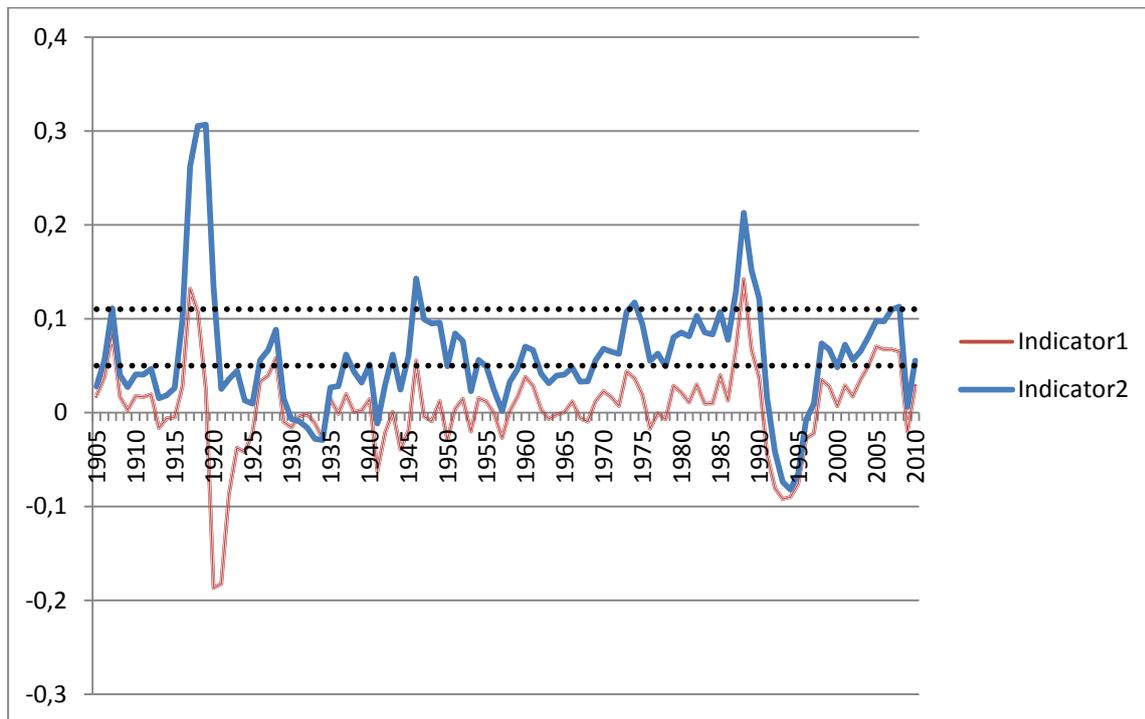
It is not obvious which values of the indicators should be regarded alarmingly high. If we apply e.g. the threshold value 0.05, the first indicator would have alarmed in 1907, 1917–1918, 1928, 1946, 1987–1989 and 2005–2008. (See chart 1) With no exception, each of these years was characterized by very strong credit growth. With the exceptions of 1907 and 1946, each alarm was followed by some kind of a financial or banking crisis. The false alarm during exceptional circumstances in 1946 may be due to rapid inflation; the moving average of the nominal GDP seemed low because the output of previous years was sold at a much lower price level. In 1907, the growth rate of bank lending was exceptionally high, but no crisis followed. Unlike in the United States, there was no financial panic in Finland in 1907.

Hence, the first indicator would have worked fairly well over this lengthy period of time. The second indicator would have performed slightly worse. Using the threshold value 0.11, the signal would have warned in 1907, 1917–1920, 1946, 1974, 1987-1990 and 2008. Hence, it would have missed the crisis of the 1930s and given an additional false alarm in 1974.

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<sup>6</sup> Thanks are due to Tapio Korhonen for collecting a large data set on the history of the Finnish economy.

Chart 1. Indicators 1 and 2 in Finland in 1905-2010



The robustness to structural changes in the economy was tested. The sample was cut into two sub-samples, 1905-1959 and 1960-2010. War years were excluded from the analysis. In the light of t-tests, there is some evidence the mean of the first indicator would have been lower during the first period, but there is no evidence the standard deviation would differ. The 2<sup>nd</sup> indicator would have had broadly similar means during the two periods, but there is some weak evidence it would have been more volatile during the first period. This, however, may not be the most meaningful way to split the sample. It may be more interesting to distinguish between the years of regulation (1936-1985) and the rest of the sample. In fact, the indicator 2 remained remarkably stable during the era of regulation. This may not imply it could not have been used; there were no serious banking crises during the era of regulation either, and the stability of the index may correctly reflect the underlying conditions rather than underestimate risks. (See appendix 6)

Hence, there is no alarmingly strong evidence to support the possible criticism that the usefulness of the proposed indicators would be restricted to one relatively short era and one particular crisis.

### 3.4 Using forecasts instead of final data

Another problem with basically any trigger indicator is the actuarial lag, the time it takes to collect all the necessary data, compile it, and to calculate provisional (let alone final) statistics to be published. The GDP for the current year is not known, and preliminary quarterly data may be published with a delay of more than two months. The most problematic variable may be borrowing by non-banks from abroad. The actuarial lag is typically somewhat shorter for loans from domestic sources, and data are available on the monthly basis.

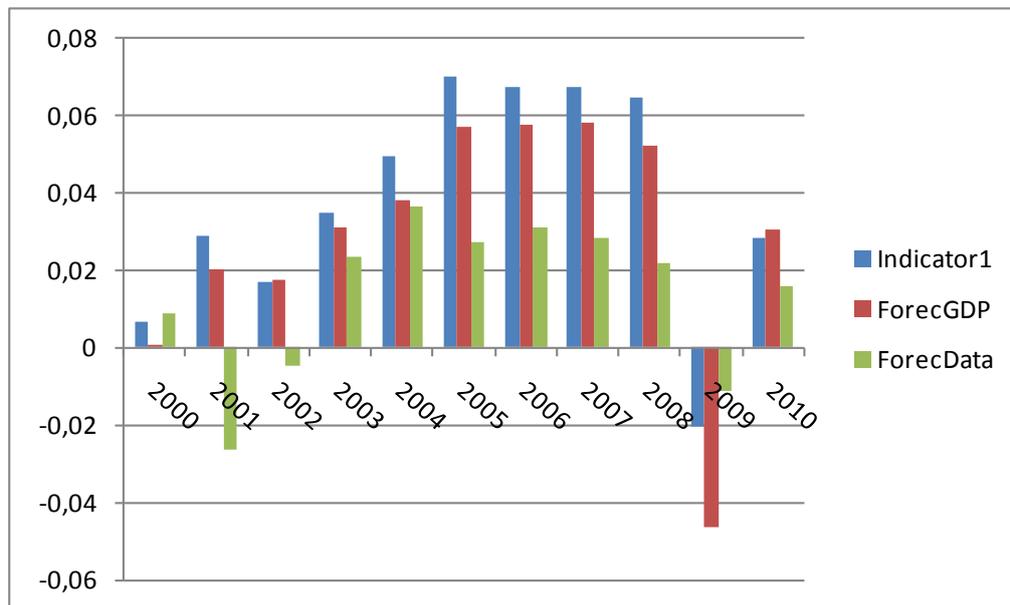
Because policy decisions should be forward looking, forecasted data might be used, and the accuracy of forecasts is highly important. In the case of the five year moving average of the nominal GDP, non-availability of GDP data for the near future may be a minor problem. The variable *ForecGDP* in chart 2 represents the first indicator when the nominal GDP of the current year is replaced by the OECD forecast published in previous December<sup>7</sup>. At least in 2000–2010, there is little difference in the results, partly because forecasts reflect future developments, and partly because the weight of the latest observation in the moving average is only 20%. The indicator based on this GDP forecast would have alarmed simultaneously with the indicator based on final data.

Moreover, the first indicator was calculated using loan stock forecasts instead of final values. As above, the GDP for the current year is replaced by the OECD forecast published in previous December. The loan forecast is based on the first Bank of Finland banking forecast of the current year. The resulting indicator (*ForecData* in Chart 2) correlates rather satisfactorily with the final value (corr coefficient +0.71 in levels), but the forecast based indicator is much less volatile. If a forecast for the loan stock is used when the trigger indicator is calculated, the threshold level for alarms should probably be much lower than 0.05.

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<sup>7</sup> Thanks are due to Maritta Paloviita for the compiled forecast data

Chart 2. . Indicator 1; calculated with final data on GDP and with data available at each moment of time



The indicator based on forecasts may be relatively reliable during tranquil times, when no policy measures would be needed, but unfortunately it probably fails when active decision making would be useful. Almost by definition, loan growth becomes excessive when it behaves in a surprising and unprecedented way. It goes without saying that every possible indicator based on these two variables suffers from the same problem irrespective of whether differences or trend deviations are calculated.

## 4 The current account and housing prices

Even though imbalances in the loan market are normally reflected in abnormal behavior of the credit-to-GDP ratio, certain other variables are also worth attention. Casual observations indicate that most crisis ridden economies have experienced a real estate boom and have been running current account deficits. Spain, Ireland, Hungary and the United States prior to 2007 are excellent examples, and so are Scandinavian countries in the late 1980s. Japan in

the 1990s may be the only example from recent decades of a banking crisis ridden developed country running a current account surplus prior to the crisis.

These casual observations are confirmed by systematic research. Several previous contributions have found that the current account has got predictive power as an early warning signal of future financial crises. According to Kaminsky and Reinhart (1999), weak exports and a resulting current account deficit are frequently observed before financial crises. In the data collected by Laeven and Valencia (2008), most financial crises occur in countries with substantial current account deficits. Barrell et al (2010) found rather strong evidence on the ability of current account deficits and housing prices to predict banking crises; when controlled for these factors, mere credit growth has no independent predictive power. However, not all the evidence is consistent with these findings; the results of Roy and Kemme (2011) on the impact of the current account are mixed. Further evidence on the ability of housing market bubbles to predict financial crises is presented by e.g. Reinhart and Rogoff (2009, p 217) .

In statistical analyses it may be difficult to differentiate between housing prices and the current account because the two variables are often highly correlated. An external deficit typically occurs when the real estate market experiences a boom, and vice versa. The housing market may even be a major driver of the current account. The wealth effect of inflated property prices boosts consumption and demand for imported goods, which affects the trade balance. At least some empirical evidence points to these kinds of effects (Roy and Kemme 2011, Fratzscher et al 2007). It is also possible that there are no causalities between these variables; the correlation is due to their joint dependence on income expectations (Calomiris & al 2009).

Kauko (2012) reached the conclusion that mere credit growth as such was no major problem at the country level in the recent international boom-bust cycle; instead, a combination of excessive credit growth and current account deficit proved much more dangerous. This finding was tested with the somewhat smaller data of 24 countries used in section 3.2. The first explanatory variable equals the 2<sup>nd</sup> indicator (formula 2) if the country was running a current account deficit in 2006, otherwise it is zero. The second explanatory variable equals the 2<sup>nd</sup> indicator if the country was running an external surplus, zero otherwise. The above mentioned result is confirmed: credit growth is a problem if and only if it occurs with a current account deficit. The current account deficit as an independent variable had no predictive power. These results must be regarded with caution because of

the rather limited number of observations, especially the limited number of countries with a current account surplus in the sample. (Detailed analyses not shown here)

If one takes a superficial look at the lengthy time series concerning Finland, the conclusion is at least weakly corroborated; extreme credit growth has been an alarming sign if and only if it has occurred simultaneously with a current account deficit and a housing boom. The rapid credit growth in 2005 – 2008 occurred during a current account surplus, and the banking sector in Finland was only moderately affected by the international crisis. High indicator values in the late 1980s occurred during an era of current account deficits. The housing price bubble of 1987-1989 was extreme (see e.g. Laakso 2000), and a very severe banking crisis followed within a few years. The first indicator reached an alarming level in 1928, when Finland was running a current account deficit of somewhat less than 6 % of GDP. Data on dwelling prices in the late 1920s is hard to find, but a boom obviously took place.<sup>8</sup> Again, a banking crisis followed in the 1930s.

The 1907 alarm was not followed by a banking crisis; no housing price data was readily available for the era, but according to Bärlund (1992, p. 42) there was a current account deficit. The alarm in 1917-1918 could be ignored because of exceptional times, even though it was followed by a banking crisis; No current account data are available, but Finland was running a foreign trade deficit.<sup>9</sup>

## 5 Conclusions and discussion

This paper has presented a few ideas and empirical tests on different possible triggers for the proposed countercyclical capital buffer. In principle, the results would be applicable to any countercyclical policy tool that requires active decision making, not only to countercyclical capital buffers.

It has been proposed that the buffer requirement should be imposed on banks if the credit-to-GDP ratio significantly exceeds its trend value. This variable may be problematic because a sharp recession makes the ratio appear high when the GDP diminishes. Moreover, a lending boom often causes a short-lived expansion of output, making the

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<sup>8</sup> According to Statistical Yearbook for 1932 (Tilastollinen vuosikirja, p 301) the number of newly constructed dwellings in cities and towns peaked in 1928 at 9547 units; in the early and mid 1920s the respective figure was less than five thousand, and in the early 1930 about two thousand.

<sup>9</sup> Statistical yearbook (Tilastollinen vuosikirja) 1918, p 219

relative loan stock appear reasonably low. This distorting effect is at least partly diluted by replacing the latest data on the GDP by its moving average over several years, which largely eliminates the impact of business cycles on the denominator.

An ideal trigger variable is stationary, or at least it must have a persistent “natural” equilibrium value, which is a slightly less restricting condition. If the actual value differs substantially from this equilibrium, there must be fundamental imbalances in the financial sector, and a correction will come sooner or later. If no such equilibrium exists, no value of the indicator is “high” or “low”.

Two different indicators calculated according to the following two principles were presented.

1. If one derives the trigger from the credit to GDP ratio, differencing the credit stock is a simpler method to make the series mean-reverting than calculating proxies for the trend deviation.
2. Instead of latest data on the GDP, one should compare the credit stock to the moving average of output. This weakens the short-term impact of changes in output on the indicator. Credit cycles are reflected in the loan stock, and the analysis should almost ignore short-term development of the GDP. This would probably be consistent with the EU draft directive published in July 2012.

Either one can calculate the difference of the loan stock and divide it by the moving average of the nominal GDP, or alternatively one can calculate the difference of credit / [moving average of GDP]. The predictive power of these indicators was tested, and the results seem promising. The H-P filter residual performed significantly weaker in predicting cross-national differences in the 2008–2009 crisis.

The idea to use credit growth as the trigger variable for countercyclical buffers is not new. Drehmann et al (2010) and Drehmann et al (2011) tested various early warning indicators, including credit growth and its deviation from the average loan growth in the past. According to the results, loan growth was a less promising candidate than the trend deviation of the credit-to-GDP ratio. However, this credit growth indicator was fundamentally different from the ones presented in this paper. Drehmann et al (2010) and Drehmann et al (2011) tested the growth of credit relative to the existing stock of loans, which may not be the best alternative. Let us assume that in the starting point a country has a credit stock of, say, 80 billion. After a period of strong credit growth the loan stock may equal 160 billion. If the loan stock still grows by an additional 10 billion, the relative growth is 6.25 %. If no growth in credit had taken place, the same 10 billion would imply a 12,5 % growth, even though in this latter

case the increase is probably less dangerous because it does not prolong a period of unsustainable loan growth.

Previous research and casual observations indicate that the financial system is more fragile if the current account is on deficit and signs of a bubble can be observed in the real estate market. This could also be taken into account in decision making. Excessive credit growth may not be an alarming sign if it is financed by collecting deposits from a thrifty domestic household sector, and if loans are used to expand the productive capacity of the corporate sector. Instead, borrowing from abroad in order to buy real estate amid a housing bubble is much more problematic from the financial stability perspective.

All the indicators are vulnerable to difficulties in forecasting future developments. Hence, shortening the actuarial lag and developing good models for forecasting loan growth in euphoric times would be at least as essential as fine tuning technical details of suggested indicators.

Internationalization may pose additional data problems. Loans granted by domestic banks are reported to domestic authorities on a monthly basis, but if companies take loans from abroad, the lag in obtaining information is somewhat longer. Information on borrowing from abroad becomes available in balance of payment statistics, and possibly other sources of data, but in most cases with a much longer delay, and provisional statistics often differ fundamentally from final ones. To make things more complicated, large corporate customers may have no home country. Loans granted to such groups may be included in the statistics of one country, but in practice be used in a different part of the world. This problem would affect any indicator based on information on the stock of loans.

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## APPENDIX 1.

Table 2. ADF tests of the credit-to-GDP ratio in levels; Nordic countries, a sample of developing countries and major advanced economies

Annual data 1986-2010

Null Hypothesis: Unit root (individual unit root process)

Modified Akaike criterion in lag length selection

Exogenous variables: individual effects

Finland, Sweden, Norway, Denmark, Iceland

Total number of observations: 121

Cross-sections included: 5

Method	Statistic	Prob.**
ADF - Fisher Chi-square	3.10019	0.979
ADF - Choi Z-stat	1.71017	0.9564

Sierra Leone, Nepal, Senegal, Niger, Nigeria

Panama, Sri Lanka

Total (balanced) observations: 175

Cross-sections included: 7

Method	Statistic	Prob.**
ADF - Fisher Chi-square	18.305	0.1932
ADF - Choi Z-stat	-0.62585	0.2657

Australia, Belgium, UK, Italy, Japan, Canada, France,

Germany, USA, Switzerland, Spain

Total number of observations: 268

Cross-sections included: 11

Method	Statistic	Prob.**
ADF - Fisher Chi-square	8.96669	0.9935
ADF - Choi Z-stat	3.67597	0.9999

Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution.

Table 3. ADF tests of differenced credit-to-GDP ratios; Nordic countries, a sample of developing countries and major advanced economies

Annual data 1986-2010

Null Hypothesis: Unit root (individual unit root process)

Modified Akaike criterion in lag length selection

Exogenous variables: individual effects

Finland, Sweden, Norway, Denmark, Iceland

Total number of observations: 121

Cross-sections included: 5	Statistic	Prob
Method		
ADF - Fisher Chi-square	29.225	0.001
ADF - Choi Z-stat	-3.18901	0.001
Hadri Z-stat	-0.9639	0.8325

Sierra Leone, Nepal, Senegal, Niger, Nigeria

Panama, Sri Lanka

Total (balanced) observations: 175

Cross-sections included: 7	Statistic	Prob
Method		
ADF - Fisher Chi-square	71.4176	0.000
ADF - Choi Z-stat	-5.37217	0.000
Hadri Z-stat	-1.00322	0.842

Australia, Belgium, UK, Italy, Japan, Canada, France,

Germany, USA, Switzerland, Spain

Total number of observations: 266

Cross-sections included: 11	Statistic	Prob
Method		
ADF - Fisher Chi-square	149.316	0.000
ADF - Choi Z-stat	-6.62553	0.000
Hadri Z-stat	-0.10039	0.540

Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution.

Null hypothesis of the Hadri test: all the series are stationary

## APPENDIX 2.

Table 4. Tendency to reversion of the credit-to-GDP ratio

Dependent Variable:  $L_t/L_{t-1}$

Method: Pooled EGLS (Cross-section random effects)

Sample: 1986 2010

Included observations: 25

Cross-sections included: 16

Total pool (unbalanced) observations: 381

Swamy and Arora estimator of component variances

White cross-section standard errors & covariance (d.f. corrected)

Variable	Coefficient	t-Statistic	Prob.
C	0.967	10.4	0.000
$L_{t-1}/L_{t-2}$	0.032	0.8	0.417
$L_{t-2}/L_{t-3}$	0.041	1.1	0.260
$L_{t-3}/L_{t-4}$	0.011	0.3	0.744
$L_{t-4}/L_{t-5}$	-0.019	-0.6	0.547

Australia, Belgium, UK, Spain, Iceland, Italy  
Japan, Canada, Norway, France, Sweden,  
Germany, Finland, Switzerland, Denmark,  
USA

### Weighted Statistics

R-squared	0.003325	Mean dependent var	1.035662
Adjusted R-squared	-0.00728	S.D. dependent var	0.126002
S.E. of regression	0.12646	Sum squared resid	6.013055
F-statistic	0.313565	Durbin-Watson stat	2.077814
Prob(F-statistic)	0.86888		

### Unweighted Statistics

R-squared	0.003325	Mean dependent var	1.035662
Sum squared resid	6.013055	Durbin-Watson stat	2.077814

## Hausman test for random fixed effects

Redundant Fixed Effects Tests  
Test cross-section fixed effects

Effects Test	Statistic	d.f.	Prob.
Cross-section F	0.741204	(15,361)	0.7422
Cross-section Chi-square	11.55695	15	0.7122

## APPENDIX 3.

Table 5. Unit root testing of indicators 1 and 2, panel data on developed countries

Null Hypothesis: Unit root (individual unit root process)

1991-2010, annual data

UK, France, Germany, Italy, Japan, USA,

Switzerland, Spain, Australia, Belgium

Sweden, Denmark, Finland, Norway, Iceland

Total number of observations: 293

INDICATOR 1	test stat	p-value
Method		
ADF - Fisher Chi-square	445.813	0.000
ADF - Choi Z-stat	-14.189	0.000
Hadri Z-stat	1.195	0.116
INDICATOR 2		
ADF - Fisher Chi-square	109.051	0.000
ADF - Choi Z-stat	-5.98665	0.000
Hadri Z-stat	0.1116	0.4556

Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution.

Null hypothesis of ADF: all the series have unit roots

Null hypothesis of the Hadri test: all the series are stationary

## APPENDIX 4.

Table 6. Explained variable: non-performing loans (NPL) in 2009 as percentage of total loans

Explained variable: Ln (non-performing loans/total loans)  
in 2009 as percentage of total loans

	eq 1	eq 2	eq 3	eq 4	eq 5	eq 6
Constant	0.858 (6.8)***	0.539 (4.3)***	0.230 (1.4)	0.241 (1.1)	0.241 (1.3)	0.529 (3.9)***
Ln(NPL/loans):s in % in 2006	0.432 (3.4)***	0.580 (5.4)***	0.598 (6.1)***	0.597 (5.6)***	0.597 (5.7)***	0.583 (5.3)***
HP-filter residual in 2006 $\lambda = 100$	0.046 (3.1)***			0.030 (0.2)	0.003 (0.2)	-0.004 (-0.2)
Indicator 1 in 2006		6.417 (5.5)**		0.010 (0.0)		6.729 (3.6)***
Indicator 2 in 2006			5.438 (6.1)***	5.227 (1.6)*	5.230 (4.1)***	
Adj R <sup>2</sup>	0.389	0.627	0.669	0.635	0.654	0.609
	eq 7	eq 8	eq 9	eq 10		
Constant	0.803 (6.4)***	0,2326 (1.0)	0.276 (1.6)	0.554 (4.2)***		
Ln(NPL/loans):s in % in 2006	0.499 (3.9)***	0,6075 (5.8)***	0.607 (5.9)***	0.584 (5.3)***		
HP-filter residual in 2006 $\lambda = 488$	0.037 (3.7)***	0.011 (0.8)	0.009 (0.8)	0.006 (0.5)		
Indicator 1 in 2006		-1.403 -0.300		5.765 (3.1)***		
Indicator 2 in 2006		5.668 (1.7)*	4.730 (4.7)***			
Adj R <sup>2</sup>	0.45	0.648	0.664	0.613		

t-values in parentheses

Cross-sectional OLS, N=24

\* = 10% significance, \*\*=5 % significance, \*\*\*=1 % significance

OLS results, cross-national data from 24 countries

## APPENDIX 5.

Table 7. ADF test statistics, Annual data 1905-2010 for Finland

### Indicator 1

Augmented Dickey-Fuller test statistic

Modified Akaike criterion in lag length selection

Variable	Coefficient	Std. Error	t-Statistic	Prob.
Indicator1(-1)	-0.366	0.076	-4.80	0.000
C	0.002	0.004	0.62	0.539

t-Statistic -4.80

Prob. (McKinnon) 0.0001 Lag Length: 0

Null Hypothesis: Indicator1 has a unit root

### Indicator 2

Augmented Dickey-Fuller test statistic

Modified Akaike criterion in lag length selection

Variable	Coefficient	Std. Error	t-Statistic	Prob.
Indicator2(-1)	-0.246	0.065	-3.81	0.000
C	0.015	0.005	2.70	0.008

t-Statistic -3.81

Prob. (McKinnon) 0.0038 Lag Length: 0

Null Hypothesis: Indicator2 has a unit root

## APPENDIX 6.

Table 8. Development of the two proposed indicators in Finland in different eras, war years excluded

		1905-1959	1960-2010
Indicator 1	Mean	-0.001	0.014
	t-value		-1.7*
	StDev	0.051	0.043
	F-ratio		1.430
		1905-1959	1960-2010
Indicator 2	Mean	0.056	0.062
	t-value		-0.459
	StDev	0.068	0.054
	F-ratio		1.61*
		1905-1934 & 1986-2010	1935-1985
Indicator 1	Mean	0.056	0.062
	t-value		-0.500
	StDev	0.068	0.054
	F-ratio		1.61*
		1905-1934 & 1986-2010	1935-1985
Indicator 2	Mean	0.058	0.060
	t-value		-0.200
	StDev	0.079	0.031
	F-ratio		6.73***

Years 1914-1919 and 1939-1946 excluded

\* = 10% significance, \*\*=5 % significance, \*\*\*=1 % significance

Finnish data from 1905–2010