Statistical Journal of the United Nations ECE 23 (2006) 119–125
IOS Press

The relationship between different price indexes: A set of evidence from inflation targeting countries

Yilmaz Akdi[a], Hakan Berument[b,*], Seyit Mümin Cilasun[c] and Hasan Olgun[d]
[a]*Department of Statistics, Faculty of Science, Ankara University 06100 Tandogan Ankara, Turkey*
[b]*Department of Economics, Bilkent University, 06800 Bilkent Ankara, Turkey*
[c]*Department of Economics, Atilim University, 06836 Incek Ankara, Turkey*
[d]*Department of Economics, Baskent University, 06530 Ankara, Turkey*

Abstract. The possible long-run relationships between the Consumer Price Index and the Wholesale Price Index are analyzed for three inflation targeting countries – Canada, Sweden and the UK – using three different statistical techniques. The Engle-Granger test finds cointegration only for Sweden. The Johansen's test and the model-free and seasonality robust periodogram based test conclusively show that the two price indexes are not cointegrated in the three countries included in the sample. Hence, the values of these indexes may consistently diverge over time. However, the two price indexes move together in the short run. These findings have some implications for the success of inflation targeting monetary policies.

Keywords: Cointegration, periodogram, price indices

1. Introduction

For the inflation or price level targeting policy goals, the selection of the basket from which the price indexes is calculated, is no less important than the level of targeted inflation. This is due to the fact that different indices can yield different inflation rates. Hence, central banks, which pursue this popular policy goal need to be very careful in choosing the relevant price index to be targeted.

The effects of central bank policies on prices can vary depending both on the price index chosen and the policy instruments used. While the exchange rate policies are more likely to affect the prices of tradable than non-tradable goods, interest rate polices have more effect on the prices of goods than services. Therefore, selection of the price index is crucial for the success and credibility of central banks.

The aim of this study is to assess the possible short-run and long-run relationships between the consumer price index (CPI) and other price indexes for three inflation targeting countries; Canada, Sweden and UK. The composition of CPI consists of both tradable and non-tradable goods/services; while, WPI includes only tradable goods in its composition. Hence, as it was mentioned before, these two indexes are affected differently by exchange rate and interest rate based policies. Therefore, it

*The corresponding author. Tel.: +90 312 290 2342; Fax: +90 312 266 5140; E-mail: berument@bilkent.edu.tr; Web: http://www.bilkent.edu.tr/~berument.

might be the case that the indices do not move together. By simply checking the cointegration of these indexes, we can decide whether their values diverge or converge in the long-run. If their values diverge overtime – that is the two series are not cointegrated, then announced inflation target might be reached for some indexes but missed for some others. This might damage the credibility of a Central Bank. The selection of the price index is therefore vital for the implementation and success of inflation target based monetary policies. However, there can be a short-run relationship between the two indices. This means their short-run behavior will not be affected by their long-run trends.

The relationship between different price indexes, to the best of our knowledge, has not been paid much attention in the literature. The relevant literature has largely focused on the effects of various economic shocks on these series and it has been found that they are affected differently by these shocks. For example, technology shocks and increasing returns to scale [16]; different stickiness of intermediate product prices [3]; tight monetary policy [5]; oil price shocks [8]; and real exchange rates [14] affect the prices of different products differently. In this study, however, rather than assessing the third variable effect on these two series, the dynamic relationship between them will be analyzed.

In particular, we look for a possible relationship between the monthly WPI and CPI series for the January 1992 – January 2005 period in three inflation targeting countries; Canada, Sweden and the UK by using conventional cointegration tests such as Engle and Granger's [9] single equation, Johansen's [13] multivariate cointegration tests as well as the periodiogram based cointegration test. The main contribution of this study is to use this third technique. This technique, which was developed by Akdi [1] and Akdi and Dickey [2] has the advantage of being model free and seasonality robust. Since the CPI and the WPI are quite seasonal series and the seasonality in the data can alter the basic inference gathered from the data ([6,12,15,17]), employing this technique will handle this problem.

The rest of the paper is organized as follows. The next section introduces the data and conventional as well as the periodogram-based testing methods, and the empirical results obtained. Section 3 searches for a possible short-run relation between the two price series. The final section concludes.

2. The data, testing techniques and empirical results

The CPI series are available for each country. However, as a proxy for the WPI we use industry selling prices (ISP) for Canada, the domestic supply prices (DSP) for Sweden and the Index of Industrial Output Prices (IOP) for the UK. The data source is International Financial Statistics (IFS) of the IMF.

As a first step we test whether each of the price series has a unit root. If a series has a unit root, then this means that a shock that hit the series will affect it forever. In this case, the series is non-stationary. On the other hand, if a series does not have a unit root (i.e., stationary), then the effect of a shock will disappear in the long-run. For this purpose the conventional Dickey-Fuller, Phillips-Peron (PP) Unit Root Tests and Dickey, Hasza and Fuller (DHF) [7] seasonal unit root tests are employed. The results for the two series are given in Table 1. Series with an intercept term are reported in Panel A, series with intercept term and the time trend are reported in Panel B and Panel C reports the tests on the first difference of the series for the Dickey-Fuller, Phillips-Peron Unit Root Tests and the twelfth difference for the Dickey, Hasza and Fuller test. According to Table 1 the null of unit root cannot be rejected in either series in levels (with and without time trend) at the 5% level, hence, the series are not stationary. However, we can reject the null of unit root at the 10% level for the UK_IOP with the ADF and PP tests, and we can reject the null of unit root at the 10% level for the UK_CPI with the intercept term and trend. However, we could reject the null of unit root in the differenced series with the PP and DHF (12) at the 1% level. When the ADF tests statistics are considered, we cannot reject the unit root for the difference

Table 1
Unit root test results

	A: Intercept			B: Intercept with Trend			C: Difference with Intercept		
	ADF	PP	DHF(12)	ADF	PP	DHF(12)	ADF	PP	DHF(12)
CA_ISP	−2.298	−2.624*	−2.481*	−1.822	−1.852	−2.100	−9.824***	−9.816***	−9.859***
CA_CPI	0.483	0.627	3.176	−2.119	−1.974	−0.710	−10.63***	−10.564***	−10.245***
SW_DSP	−1.661	−1.740	−1.190	−2.188	−1.870	−0.821	−7.350***	−7.802***	−8.905***
SW_CPI	−0.700	−2.167	0.250	−2.199	−2.515	−0.351	−3.08**	−10.534***	−7.654***
UK_IOP	−2.874*	−3.321*	−1.133	−2.643	−2.661	−2.029	−1.875	−7.957***	−7.849***
UK_CPI	−0.482	−0.884	−1.079	−3.302*	−3.180*	−0.166	−2.60*	−13.904***	−4.895***

Note: *indicates the level of significance at 10%, **indicates the level of significance at 5% and ***indicates the level of significance at 1%. The critical values are gathered from [10,11].

of UK_IOP but can reject the null at 10% level for the UK-CPI. Hence, if we opt for the 5% level of significance and the PP and DHF tests, we conclude that both series are I(1) in the three countries, that is, original series are not stationary but their first differences are stationary.

As a further test of first order integration, we apply the Akdi and Dickey's [2] periodogram based unit root test. For this test, the trigonometric transformation of the series is used. Given a time series $\{Y_1, Y_2, \ldots, Y_n\}$, the periodogram ordinate (without any model specification) is,

$$I_n(w_k) = \frac{n}{2}\left(a_k^2 + b_k^2\right) \tag{1}$$

where a_k, b_k are the Fourier coefficients and defined as,

$$a_k = \frac{2}{n}\sum_{t=1}^{n}(Y_t - \bar{Y})\cos(w_k t),\ b_k = \frac{2}{n}\sum_{t=1}^{n}(Y_t - \bar{Y})\sin(w_k t) \text{ and } \bar{Y} = \frac{1}{n}\sum_{t=1}^{n}Y_t. \tag{2}$$

When $w_k = 2\pi k/n$, the following equality obtains,

$$\sum_{t=1}^{n}\cos(w_k t) = \sum_{t=1}^{n}\sin(w_k t) = 0$$

and this causes the Fourier coefficients to be invariant to the mean and therefore the periodogram ordinate is invariant to the mean. Moreover, periodogram based unit root/cointegration tests have the advantage of being seasonality robust, and model free from the selection of the lag lengths (see [1,2]).[1]

The rejection of the null hypothesis of a unit root needs small values of the periodogram ordinates. Therefore, the values of the test statistics, $T(w_k)$ can be used to test for a unit root where,

$$T(w_k) = \frac{2(1-\cos(w_k))}{\hat{\sigma}^2} I_n(w_k). \tag{3}$$

Under the assumption of stationarity the test statistics are distributed as a mixture of chi-squares exactly for AR(1) series. In this case, the normalized periodogram will be distributed as chi-squares with two

[1] The periodogram based method has certain advantages over conventional tests. Firstly, conventional tests require the estimation of too many AR parameters to account for the dynamics/seasonality of the series. Secondly, test results change with the sample size in conventional tests, while the periodogram based method requires no parameter estimation except for variance. Thirdly, the critical values of the test statistics are free of sample size constraints. These points may constitute considerable advantages, especially for small samples.

Table 2
Periodogram based unit root test results

Series	$I_n(w_1)$	$\hat{\sigma}^2$	$T_n(w_1)$	Critical Values	Conclusion
CA_ISP (level)	0.54744	0.00004	21.91690	0.178	Unit Root
CA_CPI (level)	0.47512	0.00001	91.12120	0.178	Unit Root
CA_ISP (difference)	0.00014	0.00004	0.005894	0.178	Stationary
CA_CPI (difference)	0.00002	0.00001	0.004159	0.178	Stationary
SW_DSP (level)	0.50602	0.00004	20.2586	0.178	Unit Root
SW_CPI (level)	0.14484	0.00002	11.5974	0.178	Unit Root
SW_DSP (difference)	0.00006	0.00003	0.003568	0.178	Stationary
SW_CPI (difference)	0.000005	0.00001	0.000758	0.178	Stationary
UK_IOP (level)	0.19965	0.00001	50.9111	0.178	Unit Root
UK_CPI (level)	0.87867	0.00001	140.711	0.178	Unit Root
UK_IOP (difference)	0.00016	0.00001	0.044964	0.178	Stationary
UK_CPI (difference)	0.000005	0.00001	0.000772	0.178	Stationary

degrees of freedom asymptotically. The power of the tests is not exact in conventional tests. However, the power can be calculated analytically for the periodogram method to test for a unit root (see [1]). For higher order series, the same distribution is obtained asymptotically; that is,

$$T(w_k) = \frac{2(1-\cos(w_k))}{\hat{\sigma}^2} I_n(w_k) \xrightarrow{D} Z_1^2 + 3Z_2^2 \tag{4}$$

where Z_1 and Z_2 are independent standard normal random variables and σ^2 is the variance of the error term. Here, the notation "$\xrightarrow{D}$" stands for convergence in distribution. The critical values of this distribution are provided by Akdi and Dickey [2].

Table 2 exhibits the results for the two price series in each country. Since $T(w) > 0.178$ we fail to reject the null hypothesis of a unit root for both indexes for the three countries (i.e., these series are non-stationary). However, we can reject the null of unit root for all these series once they are differenced (i.e., differenced series are stationary). Hence, we conclude that the two price series are I(1) for the three countries (not the level but the first difference of the series are stationary). Accordingly it makes sense to search for a possible cointegration relationship between the two price series. For this purpose, we again employ three different procedures: Engle-Granger [9]; Johansen's [13]; and periodogram based cointegration tests.

The results of the Engle-Granger cointegration test are presented in Table 3. The β is the estimated value of the coefficient from regressing price series on each other. The ADF statistic for the residual term is reported in the last column. Since the value of this statistic calculated from the residuals, is greater than the critical value of -3.03 at 10% level for UK and Canada, we conclude that there is no stationary linear relationship between the price series. In other words, there is no cointegration relation between the price indexes for these two countries (i.e., these two series are not moving together in the long-run). However, we reject the null hypothesis of no cointegration for Sweden at the 10% level. Therefore, we have a long-run relation between price indexes for this country.

Table 4 reports the cointegration tests from the Johansen's method. According to the table, the null hypothesis of no cointegration cannot be rejected for any of the countries. We conclude that the two price series are not cointegrated in the three countries.

Table 5 reports the periodogram based cointegration test results. The $\hat{\beta}$ in the table is the estimated slope of the regression of the real part of the cross periodogram (WPI and CPI) on the periodogram of the CPI. τ_a is the value of the t-statistic obtained from periodogram based regression. The critical value

Table 3
Engle-Granger cointegration test results

Cointegration relation	β	ADF statistics of residuals
Canada	1.048320	−1.884615
Sweden	1.637876	−3.146643*
UK	0.526688	−2.375105

*Indicates the level of significance at 10%.

Table 4
Johansen's cointegration tests results

Cointegration Relation	Hypothesized No. of CE(s)	Eigenvalue	λ-Trace Statistic	λ max Statistic	λ trace test At 5%	λ max test At 5%
CA (isp − cpi)	None*	0.074165	11.66982	11.55894	15.41	14.07
	At most 1*	0.000739	0.110880	0.110880	3.76	3.76
SW (dsp − cpi)	None*	0.056326	14.17005	8.696121	15.41	14.07
	At most 1*	0.035835	5.473927	5.473927	3.76	3.76
UK (iop − cpi)	None*	0.079001	12.94735	12.34437	15.41	14.07
	At most 1*	0.004012	0.602976	0.602976	3.76	3.76

Table 5
Periodiogram based cointegration test results

Relation	$\hat{\beta}$	τ_a	Critical Value %5	Conclusion
Canada	0.772963	−0.656	−3.43564	No Cointegration
Sweden	0.545901	−0.212	−3.43564	No Cointegration
UK	1.794085	−2.209	−3.43564	No Cointegration

The critical values for the test statistics are taken from [4].

at 5% level of significance has been constructed by simulation with 20 000 replications.[2] The results clearly show that there is no cointegration between the WPI and CPI in the three countries.

3. Short-run relations

The former section has refuted any cointegration relation between the two price series in the three countries. Yet it is legitimate to ask whether a short-run relation exists. More specifically we asked if there is a one-to-one relationship between CPI and WPI inflations. For this purpose we regress the first difference of log of WPI on the first difference of log of CPI. The results presented in panel A of Table 6 show that there is a strong and statistically significant short-run relation between the two price series in the three countries. Panel B of the table repeats the same exercise by including dummy variables to account for seasonality. Panel B confirms the findings of panel A. However, we reject that this correlation is one-to-one. That is we reject the null-hypothesis that the coefficient of the logarithmic first difference of the WPI is one. Therefore we conclude that a monthly change in CPI leads to a monthly change in WPI in the same direction but not by the same amount. The policy implication of this finding is that for a given set of monetary and foreign exchange rate policies measured inflation will vary depending on the price index.

[2] For more details see [1]. For the construction of the critical values see [4].

Table 6
Short-run relationship between price indexes

	Panel A: Without Dummies	Panel B: With Dummies
Canada	0.127639	0.125464
	(3.674229)	(3.602747)
Sweden	0.235175	0.205713
	(4.2388)	(4.621348)
UK	0.394667	0.270886
	(3.562603)	(4.250685)

4. Conclusions

The objective of this paper was to determine the nature of the long run relationship between CPI and WPI series. Accordingly, we have searched for cointegration between the price series. The Johansen's test and the periodogram-based test have shown that they are not cointegrated in any country. However, the Engle-Granger test has detected a possible cointegration between the CPI and WPI in Sweden but not for UK and Canada. Thus, we can claim that the two price indexes may consistently diverge over time in the three countries. Hence, the announced inflation target might be reached for some indexes but missed for some others. At this stage of our research we can only suggest that instead of a single price index it might be advisable for central banks to target a weighted average of two or more price indexes.

References

[1] Y. Akdi, *Periodogram Analysis for Unit Roots,* Ph.D. Dissertation, NCSU, 1995.
[2] Y. Akdi and D.A. Dickey, Periodograms for unit root time series: Distributions and tests, *Communications in Statistics* **27**(1) (1998), 69–87.
[3] S. Basu, Intermediate goods and business cycles: Implications for productivity and welfare, *American Economic Review* **85** (1995), 512–531.
[4] H. Berument, Y. Akdi and C. Atakan, An empirical analysis of Istanbul Stock Exchange sub-indexes, *Studies in Nonlinear Dynamics & Econometrics* **9**(3) (2005), Article 5.
[5] T.E. Clark, The response of prices at different stages of production to monetary policy shocks, *The Review of Economics and Statistics* **81**(3) (1999), 420–433.
[6] G. Cubadda, Common cycles in seasonal non-stationary time series, *Journal of Applied Econometrics* **14** (1999), 273–291.
[7] D.A. Dickey, H.P.Hasza and W.A. Fuller, Testing for unit roots in seasonal time series, *Journal of the Americal Statistical Association* **79** (1984), 355–367.
[8] K. Doroodian and R. Boyd, The linkage between oil price shocks and economic growth with inflation in the presence of technological advances: a CGE model, *Energy Policy* **31**(10) (2003), 989–1006.
[9] R.F. Engle and C.W.J. Granger, Cointegration and error correction: Representation, estimation, and testing, *Econometrica* **55** (1987), 251–276.
[10] P.H. Franses and B. Hobijn, Critical values for unit root tests in seasonal time series, *Journal of Applied Statistics* **24**(1) (1997), 25–47.
[11] J.D. Hamilton, *Time Series Analysis,* Princeton, NJ, Princeton University Press, 1994.
[12] A. Hecq, Does seasonal adjustment induce common cycles? *Economic Letters* **59** (1998), 289–297.
[13] S. Johansen, Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control* **12** (1988), 231–254.
[14] J. Kim, Short run real exchange rate dynamics: A SUR approach, *Applied Economics Letters* **11** (2004), 909–913.
[15] A. Maravall, Unobserved components in econometric time series, in: *Handbook of Applied Econometrics,* (Vol. 1), M.H. Pesaran and M.R. Wickens, eds, Blackwell, Massachusetts, 1995, pp. 12–72.
[16] K.M. Murphy, A. Shleifer and R.W. Vishny, Building blocks of market clearing business cycle models, *NBER Macroeconomics Annual* (1989), 247–286.
[17] Y.-W. Cheung and F. Westermann, Sectoral trends and cycles in Germany, *Empirical Economics* **28**(1) (2003), 141–156.

Hakan Berument is an Associate Professor in the Department of Economics at Bilkent University. He has received his Ph.D. from the University of North Carolina in 1994. He has also served the director of the Research Department at the Central Bank

of the Republic of Turkey. His research interests include sports economics, effectiveness of monetary policy and international transmission of international shocks.

Yılmaz Akdi is an Associate Professor in the Department of Statistics at Ankara University. He has received his Ph.D. from North Carolina State University in 1995. His research interests include time series and linear models.

Seyit Mümin Cilasun is a research assistant in the Department of Economics at Atılım University and a Ph.D. student in the Middle East Technical University. His research interests are industrial organization, microeconomics and applied econometrics.

Hasan Olgun is a professor at the Department of Economics, Baskent University-Ankara. He has received his Ph.D from the Department of Political Economy, Johns Hopkins University in 1973. He has participated as a country author in the Agricultural Pricing Policies of the World Bank, directed by Anne Krueger. His current research interest covers international economics, monetary economics and macroeconomics.