香 港 樹 仁 大 學

Evidence from asymmetric adjustment and causality between consumer and producer prices in the four selected European economies

Shu-kam LEE
Cho-yiu Joe NG
Kai-yin WOO

August 2014

經濟及金融學系

Working Paper Series

Department of Economics and Finance

Hong Kong Shue Yan University

Working Paper Series
August 2014

ISBN:
The URL is:
http://www.hksyu.edu/economics/working_paper/2014/Working%20paper_2014_Aug_Lee.pdf

Information on the Working Paper Series can be found on the last page. Please address any comments and further inquiries to:

Dr. Yi-Man Rita Li
Working Paper Coordinator
Department of Economics and Finance
Hong Kong Shue Yan University
10 Wai Tsui Crescent
Braemar Hill Road
North Point
Hong Kong
Fax: 2806-8044
Tel: 2806-5179 (Mr. Ken Tsui)
Email: ymli@hksyu.edu

Evidence from asymmetric adjustment and causality between consumer and producer prices in the four selected European economies

Shu-kam LEE
Department of Economics and Finance
Hong Kong Shue Yan University

Cho-yiu Joe NG
Department of Economics and Finance,
City University of Hong Kong

Kai-yin WOO
Department of Economics and Finance
Hong Kong Shue Yan University

Abstract

This paper examines Granger-causality between consumer price index (CPI) and producer price index (PPI) in Germany, United Kingdom, France and Italy, based on a Momentum Threshold Vector Error Correction Model (M-TVECM) with asymmetric adjustment process. Our study introduces theoretical ground of price rigidity to the CPI-PPI relation and allows for threshold cointegration. The findings indicate CPI and PPI is cointegrated in an asymmetric manner in all countries. The M-TVECM suggests, for Germany and Italy, producer prices are relatively sticky downward in response to a decrease in consumer prices than increase. Consumer prices are sticky in response to either an increase or a decrease in producer prices in all countries except for United Kingdom that shows upward rigidity of consumer prices. For France and United Kingdom, we find evidence of sticky downward producer prices in response to a decrease in consumer prices. Causality test indicates 'production chain view' is relevant only in the short run, and 'demand-pull view' is relatively relevant in the long run for the four countries.

JEL classification codes: C12; C22; E31, O52

Keywords: Threshold cointegration, Granger causality; asymmetric adjustment; Europe

I. Introduction

The causal relationship between Consumer Prices Index, CPI, and Producer Prices Index, PPI, has been a long debate for centuries. Particular attention has been paid because identifying the causality can help central banks to improve the forecast performance of inflation and define their inflation targets (among others, see Tiwari, 2012; Tiwari et al., 2014). There are basically two theoretical points of view, namely 'production chain view' and 'demand-pull view', justifying different directions of causality between CPI and PPI. The 'production chain view' suggests that through the production chain, shocks to raw material prices will eventually transmit to prices of intermediate goods as well as final goods sold to the consumers. Hence, PPI leads CPI. As reported by Cushing and Mcgarvey (1990), early literature like Engle (1978), Silver and Wallace (1980) and Guthrie (1981) is based on a one-sided distributed lag model with producer prices leading. The 'demand-pull view' suggests an opposite direction of causality based on the derived demand analysis (among others, see Colclough and Lange (1982); Granger et al. (1986); Jones (1986)). Demand for a factor of production depends on the demand for the final goods and hence, CPI leads PPI.

Unfortunately, the causality between CPI and PPI was so far inconclusive yet. As criticized by Clark (1995), the pass-through mechanism of the production chain may be weak since crude materials, intermediate goods, and final goods may not follow a simple production chain. Some recent empirical studies, like Tiwari (2012), Akdi et al. (2006a) and Akdi et al. (2006b) could only find a uni-directional causality running from CPI to PPI. However, other recent studies, including Caporale et al. (2002), Akçay (2011) and Tiwari et al. (2014), find supports for causality with producer prices leading.

Yet, many of these studies neither attempt nor success to detect a cointegration relationship among CPI and PPI. One possible reason is that the underlying cointegration relationship is not a linear one. Moreover, previous studies may suffer from misspecification of dynamics. The possibility of asymmetric adjustment process could be brought by the Sticky Price Theory. Due to, for example, downward price rigidity, a positive shock in PPI may cause CPI to go up but a negative shock in PPI may not cause CPI to go down. From this theoretical point of view, ignoring price rigidity may cause serious bias in finding the causality between CPI and PPI in a symmetric adjustment framework.

Retailers may not change their prices frequently due to menu cost. Even a small amount, menu cost may be sufficient to generate aggregate nominal rigidity (see, among others, Caplin and Leahy (1991, 1997); Caplin (1993); Levy et al. (1997) and the references therein). For both wholesale and retail markets, oligopolies who fear of price war also have little incentive to change prices. In addition to the traditional reasons for price rigidity, Borenstein and Shepard (2002) report that firms tend to spread price adjustment over several periods to minimize the cost of production and distribution causing gasoline prices to adjust incompletely to crude oil price shocks. For factor market, Dibooglu and Enders (2002) find new evidence for downward real wage rigidity in Canada.

Given the theoretical ground of price rigidity, this paper contributes to literature by considering asymmetric adjustment process between CPI and PPI. We employ the threshold cointegration test developed by Enders and Siklos (2001) and reexamine the causality between CPI and PPI in the four biggest economies in Europe (in term of purchasing power parity, PPP, GDP in 2013), namely Germany, United Kingdom, France and Italy, based on a Momentum Threshold Vector Error Correction Model, M-TVECM. Since M-TVECM is used, we are able to report both short-run and long-run causality[1].

The rest of this paper is organized as follows: Methodology and data will be presented in next section, followed by the results and conclusion.

II. Methodology and Data

Following the standard testing procedure, the first step is to conduct unit root test in order to avoid spurious regression problem (Granger and Newbold, 1974). We employ the Dickey Fuller generalized least squares (GLS) test (Elliott, et al., 1996) to test for the stationarity of the CPI and PPI series. The lag length selection is based on Bayesian information criterion (BIC). If the variables are I(1), i.e. each of them contains a unit root at level but not at 1st difference form, we then proceed to cointegration test which confirms the long-run equilibrium relationship between CPI and PPI and estimate the cointegrating equation using ordinary least squares (OLS) as:

[1] Tiwari (2012) also reports short-run and long-run causality between CPI and PPI based on Squared Wavelet Coherence (WTC) approach.

$$\text{Equation 1: } CPI_t = \beta_0 + \beta_1 PPI_t + \mu_t$$

where μ_t is the residual.

The Engle-Granger cointegration test is to test if the residual in equation 1, μ_t in is stationary. Using the residual, μ_t, we may estimate the following Autoregression model:

$$\text{Equation 2: } \Delta\mu_t = \rho\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-i} + \varepsilon_t$$

Note that Equation 2 does not account for asymmetric adjustment. With the theoretical ground of price rigidity, it is not appropriate to apply traditional linear cointegration test. Pippenger and Goering (1993) and Enders and Granger (1998) point out that the standard linear cointegration test has low power when facing misspecification of dynamics. As suggested by Enders and Siklos (2001), Equation 2 can be modified to introduce asymmetric adjustment by letting the discrepancies from the long-run equilibrium, i.e. μ_t, to behave as a Threshold Autoregressive (TAR) process and a Momentum-Threshold Autoregressive (M-TAR) process. As central banks always forecast inflation rates and define their inflation targets, we assume the market agents do not care much about the index level. In this paper, we apply only the M-TAR process since we focus upon the inflation rate of consumers' and producers' goods[2].

Considering the M-TAR modification to Equation 2 as follows:

$$\text{Equation 3: } \Delta\mu_t = M_t\rho_1\mu_{t-1} + (1 - M_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-i} + \varepsilon_t$$

where M_t is the Heaviside indicator function, that is:

$$\text{Equation 4: } M_t = \begin{cases} 1 & if\ \Delta\mu_{t-1} \geq \tau \\ 0 & if\ \Delta\mu_{t-1} < \tau \end{cases}$$

where τ is the threshold parameter.

Since we allow for an unknown threshold in the above M-TAR model, Chan (1993) method is employed to find the super-consistent estimate of the threshold parameter. We order the threshold variable (i.e. $\Delta\mu_{t-1}$) from the lowest to the highest

---

[2] In TAR model, Equation 4 becomes $M_t = \begin{cases} 1 & if\ \mu_{t-1} \geq \tau \\ 0 & if\ \mu_{t-1} < \tau \end{cases}$. Therefore, TAR model defines threshold parameter as level of μ_{t-1} and it focuses on index level but not the first difference of price index, namely inflation rate.

value. It is a usual practice to exclude the highest and lowest 15% of the $\Delta\mu_{t-1}$ from the grid search so as to ensure an adequate number of observations on each side of the threshold. Then, we estimate the M-TAR model in Equations 3 and 4 using each value of $\Delta\mu_{t-1}$ as a threshold. The value of $\Delta\mu_{t-1}$ resulting in the lowest residual sum of squares is the super-consistent estimate of the threshold parameter.

It is possible to test for an attracter although the adjustment process is nonlinear. A sufficient condition for μ_t to be stationary is $-2 < (\rho_1, \rho_2) < 0$ for any given value of τ (Divooglu and Enders, 2001).[3] Moreover, Tong (1983; 1990) showed that if μ_t is stationary, the least square estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution. If we can reject the null hypothesis of $\rho_1 = \rho_2 = 0$, than it can be concluded that there is an attracter, i.e. μ_t is stationary and a cointegrating relationship exists among the variables in Equation 1. We can calculate F-statistic for the null hypothesis of $\rho_1 = \rho_2 = 0$ but, since the F-statistic does not follow a standard distribution, it is inappropriate to compare this sample statistic, called Φ_μ statistic, with the standard critical values of F-distribution. Instead, we simulate the critical values of Φ_μ by using Monte Carlo simulations.

Having concluded that the variables are cointegrated, next we test for symmetric against asymmetric adjustment. Since the asymptotic joint distribution of ρ_1 and ρ_2 converges to a multivariate normal, the null hypothesis of $\rho_1 = \rho_2$ (i.e. symmetric adjustment) can be tested by using a standard F-test (Enders and Siklos, 2001). If the null hypothesis of $\rho_1 = \rho_2$ is rejected, we can conclude CPI and PPI are cointegrated with asymmetric adjustment based on M-TAR process.

Finally, we estimate M-TVECM system by using full information likelihood as follows:

Equation 5:

$$\Delta CPI_t = M_t\rho_{11}\mu_{t-1} + (1 - M_t)\rho_{12}\mu_{t-1} + \sum_{i=0}^{p}\alpha_{1i}\Delta PPI_{t-i} + \sum_{j=0}^{p}\alpha_{2j}\Delta CPI_{t-i} + u_t$$

$$\Delta PPI_t = M_t\rho_{21}\mu_{t-1} + (1 - M_t)\rho_{22}\mu_{t-1} + \sum_{i=0}^{p}\alpha_{2i}\Delta PPI_{t-i} + \sum_{j=0}^{p}\alpha_{2j}\Delta CPI_{t-i} + v_t$$

where M_t is the M-TAR Heaviside indicator function given by Equation 4. Different adjustment speeds are incorporated in the M-TVECM, represented by ρ_{11} and ρ_{21} (above the threshold), and, ρ_{12} and ρ_{22} (below the threshold). Long-run causality can be found by testing the joint significance for the adjustment coefficients of error

[3] Petrucelli and Woolford (1984) stated a weaker set of sufficient conditions for the stationarity of μ_t is $\rho_1 < 0$, $\rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2) < 1$ for any given value of τ.

terms in each equation (i.e. ρ_{11} and ρ_{12}; ρ_{21} and ρ_{22}). Similarly, short-run causality can be detected by testing the joint significance for the coefficients of lagged change of explanatory variables in each equation (i.e. $\alpha_{11}= \alpha_{12} \ldots. = \alpha_{1P} = 0$; $\alpha_{21}= \alpha_{22} \ldots = \alpha_{2P} = 0$)

We collect monthly data of PPI and CPI for Germany, United Kingdom, France and Italy from Federal Reserve Bank of St. Louis. Since PPI data for France is available from January 1995, the sample period runs January 1995 to December 2013, with 228 observations in total, for each country. All variables are in natural logarithms and seasonally adjusted using the X12 procedure.

III. Result

The results of the Dickey-Fuller generalized least squares (DF-GLS) test are reported in Table 1. It cannot reject the null of existence of a unit root at level but it rejects the null at 1st difference form for all the series. We conclude that all the series are I(1).

Then, we process to the threshold cointegration test based on M-TAR model. Table 2 shows the estimated long-run relation. All the coefficients, including the constant β_0 and the slope coefficients β_1, are significant at 1% level. Threshold cointegration results are presented in Table 3. Lag length selection is based on BIC. The Φ_μ statistic rejects the null hypothesis of $\rho_1 = \rho_2 = 0$ and indicates that cointegrating relationship exists between CPI and PPI for all countries. Evidence of asymmetric adjustment is also supported by F-test, which strongly rejects the null of symmetric adjustment $\rho_1 = \rho_2$ at 1% level of significance for all countries. The threshold parameter τ in each country is approximately zero. Notice that in each case, positive deviations from the threshold are eliminated faster than negative deviations. Note that the full picture of adjustment process cannot be revealed in the coefficients of cointegration relations. The M-TVECM is more informative.

Table1. DF-GLS Unit Root Test

Country	CPI				PPI			
	level	Lag	1st difference	lag	Level	Lag	1st difference	lag
Germany	-1.8839	0	-17.0525***	0	-2.2845	3	-3.9721***	2
United Kingdom	0.0405	1	-7.3910***	1	-0.8880	1	-10.0329***	0

France	-1.4446	0	-13.1995***	0	-1.5404	1	-8.9932***	0
Italy	-0.8047	2	-3.2193**	1	-2.0913	1	-4.3474***	1

Notes: A constant and a linear trend are included in the unit root test.

Lag length selection is based on BIC.

(***) and (**) denotes the statistical significance at the 1% and 5% level, respectively.

Table 2. Estimated Long-Run Relation

Country	Cointegrating Equation	
	$\boldsymbol{\beta_0}$	$\boldsymbol{\beta_1}$
Germany	1.0300***	0.7737***
United Kingdom	2.2762***	0.5063***
France	0.4355***	0.9010***
Italy	0.8779***	0.8078***

Notes: (***) denotes the statistical significance at the 1%.

Cointegrating Equation is in the following form: $CPI_t = \beta_0 + \beta_1 PPI_t + \mu_t$

Table 3. Threshold Cointegration Test Based on M-TAR Process

Country	Lags	$\boldsymbol{\rho_1}$	$\boldsymbol{\rho_2}$	$\boldsymbol{\Phi_\mu}$	$\boldsymbol{\rho_1 = \rho_2}$	Threshold τ
Germany	3	-0.1020***	-0.0283**	9.8715	5.9179***	0.0022
United Kingdom	5	-0.0698***	-0.0098	7.0565	7.5130***	0.0028
France	3	-0.0669***	-0.0151	9.3919	5.7753***	0.0005
Italy	3	-0.0626***	-0.0161*	10.0815	6.9371***	0.0026

Notes: (***), (**) and (*) denotes the statistical significance at the 1%, 5% and 10% level, respectively.

Lag length selection is based on BIC.

The M-TAR model is in the form of Equations 3 and 4.

Using the estimated long-run relation in Table 2, we estimate M-TVECM system by using full information likelihood as reported in Table 4. System Residual Portmanteau Tests for Autocorrelations is used to determine the lag length in M-TVECM system where lag length is selected by rejecting the null of no residual autocorrelations to (at least) 20 lags. Notice that the threshold parameter τ in each country is approximately zero as reported in Table 3, we assume, for simplicity, it is zero to explain the results.

Germany and Italy show similar error correction pattern in which an increase in CPI causes PPI at a quicker pace than decrease. With any disequilibrium resulting from an increase in CPI at time t-1, PPI will adjust upward by 9.12% for Germany and 8.48% for Italy in the next month. However, with any disequilibrium resulting from a decrease in CPI at time t-1, PPI will adjust downward by only 3.98% for Germany and 2.29% for Italy in the next month. For both countries, CPI does not response to any disequilibrium resulting from an increase or a decrease in PPI since ρ_{11} and ρ_{12} are insignificant. These results imply that, in Germany and Italy, producer prices are relatively sticky downward in response to a decrease in consumer prices than increase, and consumer prices are sticky in response to either an increase or a decrease in producer prices.

The error correction pattern for France is very close to the two former countries, except ρ_{22} is insignificant. The results indicate that with any disequilibrium resulting from an increase in CPI at time t-1, PPI will adjust upward by 10.23% in the next month. However, PPI will not adjust in response to any disequilibrium resulted from a decrease in CPI at time t-1. Here, we find evidence of sticky downward producer prices. Similar to Germany and Italy, consumer prices are sticky in response to either an increase or a decrease in producer prices in France.

For United Kingdom, we also find evidence of sticky downward producer prices in response to a decrease in consumer prices. As indicated by the results that ρ_{12} is significant at 5% level and ρ_{22} is insignificant, PPI will only adjust upward by 10.56% in the next month in response to any disequilibrium caused by an increase in CPI at time t-1, while PPI will not adjust downward in response to any disequilibrium resulted from a decrease in CPI at time t-1. On the other hand, the adjustment process of CPI in United Kingdom is quite different from other countries. ρ_{12} is significant at 5% level and this shows that CPI will adjust downward by 1.42% in response to any disequilibrium caused by a decrease in PPI. However, ρ_{11} is significant which implies that CPI will not adjust in response to any disequilibrium caused by increase in PPI. This indicates consumer prices are sticky upward in response to an increase in producer prices.

Table 4. Momentum Threshold Vector Error Correction Model (M-TVECM)

Coefficient	**Country**							
	Germany		United Kingdom		France		Italy	
	Dependent variable							
	ΔCPI_t	ΔPPI_t	ΔCPI_t	ΔPPI_t	ΔCPI_t	ΔPPI_t	ΔCPI_t	ΔPPI_t

ρ_{11}	-0.0238		-0.0041		0.0073		0.0053	
ρ_{12}	-0.0016		-0.0142**		-0.0023		-0.0038	
ρ_{21}		0.0913**		0.1056**		0.1023***		0.0848**
ρ_{22}		0.0398**		-0.0124		-0.0051		0.0229*
Lags (p)	3		1		8		4	

Notes: (***), (**) and (*) denotes the statistical significance at the 1%, 5% and 10% level, respectively.

M-TVECM is in the following form:

$$\Delta CPI_t = M_t\rho_{11}\mu_{t-1} + (1 - M_t)\rho_{12}\mu_{t-1} + \sum_{i=0}^{p}\alpha_{1i}\Delta PPI_{t-i} + \sum_{j=0}^{p}\alpha_{2j}\Delta CPI_{t-i} + u_t$$

$$\Delta PPI_t = M_t\rho_{21}\mu_{t-1} + (1 - M_t)\rho_{22}\mu_{t-1} + \sum_{i=0}^{p}\alpha_{2i}\Delta PPI_{t-i} + \sum_{j=0}^{p}\alpha_{2j}\Delta CPI_{t-i} + v_t$$

Table 5. Granger Causality Test Based on M-TVECM

Country	**Hull Hypothesis**	**Wald Statistics**	**Direction of Causality**
Short- run causality			
Germany	PPI does not Granger cause CPI	8.8494**	CPI⇆PPI
	CPI does not Granger cause PPI	11.2109***	
United Kingdom	PPI does not Granger cause CPI	24.5411***	PPI→CPI
	CPI does not Granger cause PPI	0.8309	
France	PPI does not Granger cause CPI	18.8233**	PPI→CPI
	CPI does not Granger cause PPI	4.3291	
Italy	PPI does not Granger cause CPI	10.8671**	PPI→CPI
	CPI does not Granger cause PPI	0.3910	
Long-run causality			
Germany	PPI does not Granger cause CPI	1.8787	CPI→PPI
	CPI does not Granger cause PPI	13.8001***	
United Kingdom	PPI does not Granger cause CPI	4.1327	CPI→PPI
	CPI does not Granger cause PPI	9.7389***	
France	PPI does not Granger cause CPI	0.4131	CPI→PPI
	CPI does not Granger cause PPI	15.8885***	
Italy	PPI does not Granger cause CPI	1.4729	CPI→PPI
	CPI does not Granger cause PPI	12.9056***	

Notes: (***), (**) and (*) denotes the statistical significance at the 1%, 5% and 10% level, respectively.

Different from previous literature, we are able to find cointegrating relationship among CPI and PPI based on threshold cointegration test. Therefore, short-run and long-run causality test can be done based on M-TVECM. In Table 5, this is clear that long-run causality runs from CPI to PPI in each country. In the short-run, causality runs from PPI to CPI in United Kingdom, France and Italy while a bi-directional causality can be found in Germany. Generally, we conclude 'production chain view' is relevant only in short-run and 'demand-pull view' is relatively relevant in long-run for the four countries.

IV. Conclusion

This paper introduces theoretical ground of price rigidity to the CPI-PPI relation and allows for threshold cointegration developed by Enders and Siklos (2001). We examine Granger-causality between consumer and producer prices in Germany, United Kingdom, France and Italy, based on a Momentum Threshold Vector Error Correction Model (M-TVECM). The findings indicate CPI and PPI are cointegrated in an asymmetric manner in all countries. The M-TVECM suggests, for Germany and Italy, producer prices are relatively sticky downward in response to a decrease in consumer prices than increase. Consumer prices are sticky in response to either an increase or a decrease in producer prices in all countries except for United Kingdom which shows upward rigidity of consumer prices. For France and United Kingdom, we find evidence of sticky downward producer prices in response to a decrease in consumer prices. Causality test indicates 'production chain view' is relevant only in the short-run and 'demand-pull view' is relatively relevant in the long-run for the four countries.

Reference

Akçay, S. (2011) The Causal Relationship between Producer Price Index and Consumer Price Index: Empirical Evidence from Selected European Countries, International Journal of Economics and Finance, **3**, 277–32.

Akdi, Y., Berument, H. and Cilasun, S.Y. (2006a) The relationship between different price indices: evidence from Turkey, Physica A, **360**, 483–92.

Akdi, Y., Berument, H., Cilasun, S.M. and Olgun, H. (2006b) The relationship between different price indexes: a set of evidence from inflation targeting countries, Stat. J. Unit. Nation. Econ. Comm. Eur., **23**, 119–25.

Borenstein, S. and Shepard, S. (2002) Sticky prices, inventories, and market power in wholesale gasoline markets, RAND Journal of Economics, **33**, 116–39.

Caplin, A. S. (1993) Individual inertia and aggregate dynamics, in Optimal Pricing, Inflation, and the Cost of Price Adjustment, (Eds) E. Sheshinski and Y. Weiss, The MIT Press, Cambridge.

Caplin, A. S. and Leahy J. (1991) State dependent pricing and the dynamics of Money and Output, Quarterly Journal of Economics, **106**, 683–708.

Caplin, A. S. and Leahy J. (1997) Aggregation and optimisation with state-dependent pricing, Econometrica, **65**, 501–25.

Caporale, G. M., Katsimi, M. and Pittis, N. (2002) Causality links between consumer and producer prices: some empirical evidence, Southern Economic Journal, **68**, 703–11.

Chan, K. S. (1993) Percentage Points of Likelihood Ratio Tests for Threshold Autoregression, Journal of the Royal Statistical Society, 53, 691–96.

Clark, T. E. (1995) Do Producer Prices Lead Consumer Prices?, Federal Reserve Bank of Kansas City Economic Review, **80**, 25–39.

Colclough, W. G. and Lange, M. D. (1982) Empirical evidence of causality from consumer to wholesale prices, Journal of Econometrics, **19**, 379–84.

Cushing, M. J. and Mcgarvey, M. G. (1990) Feedback between wholesale and consumer price inflation: a reexamination of the evidence, Southern Economic Journal, **56**, 1059–72.

Dibooglu, S. and Enders, W. (2002) Do Real Wages Respond Asymmetrically to Unemployment Shocks? Evidence from the U.S. and Canada, Journal of Macroeconomics, **23**, 495–515.

Elliott, G., Rothenberg, T., and Stock, J. (1996) Efficient tests for an autoregressive unit root, Econometrica, **64**, 813–36.

Enders, W. and Granger, C.W.J. (1998) Unit-Root Tests and Asymmetric Adjustment With an Example Using the Term Structure of Interest Rates, Journal of Business and Economic Statistics, **16**, 304–11.

Enders, W. and Siklos, P. L. (2001) Cointegration and Threshold Adjustment, Journal of Business and Economic Statistics, **19**, 166–76.

Engle, R. F. (1978) Testing price equations for stability across spectral frequency bands, Econometrica, **46**, 869–81.

Granger, C. W. J. and Newbold, P. (1974) Spurious regressions in econometrics, Journal of Econometrics, 2, 111–20.

Granger, C. W. J., Robins, R. P. and Engle, R. F. (1986) Wholesale and retail prices : bivariate time-series modeling with forecastable error variances, in Model Reliability, (Eds) D. A. Belsley and E. Kuh, The MIT Press, Cambridge, pp. 1–17.

Guthrie, R. S. (1981) The relationship between wholesale and consumer prices, Southern Economic Journal, **47**, 1046–55.

Jones, J. D. (1986) Consumer Prices, Wholesale Prices, and Causality, Empirical Economics, **11**, 41–55.

Levy, D., Bergen, M., Dutta, S. and Venable, R. (1997) The magnitude of menu cost: direct evidence from large U.S. supermarket chains, Quarterly Journal of Economics, **112**, 791–824.

Silver, J. L. and Wallace, T. D. (1980) The lag relationship between wholesale and consumer prices, Journal of Econometrics, **12**, 375–87.

Tiwari, A. K. (2012) An empirical investigation of causality between producers' price and consumers'price indices in Australia in frequency domain, Economic Modelling, **29**, 1571–78.

Tiwari, A. K., Suresh, K.G., Arouri, M. and Teulon, F. (2014) Causality between consumer price and producer price: Evidence from Mexico, Economic Modelling, **36**, 432–40.

Tong, H. (1983) *Threshold Models in Nonlinear Time Series Analysis*, Springer, New York.

Tong, H. (1990) *Nonlinear Time Series: A Dynamical System Approach*, Oxford University Press, Oxford.

Petrucelli, J. D. and Woolford, S. W. (1984) A threshold AR(1) model, Journal of Applied Probability, **21**, 270–86.

Pippenger, M. K., and Goering, G. E. (1993) A note on the empirical power of unit root tests under threshold processes, Oxford Bulletin of Economics and Statistics, **55**, 473–81.

The working paper series is a series of occasional papers funded by the Research and Staff Development Committee. The objective of the series is to arouse intellectual curiosity and encourage research activities. The expected readership will include colleagues within Hong Kong Shue Yan University, as well as academics and professionals in Hong Kong and beyond.

Important Note

Department of Economics and Finance

Hong Kong Shue Yan University