Price Convergence and Monetary Policy in United States Cities: A Cointegration Approach with Two Structural Breaks

Rangkakulnuwat P., Morshed AKM M., Wang, H. H. and Ahn, S. K

Abstract— **This paper investigates how the price indexes of the United States 12 cities response to the shock from a city and from monetary policy. We found that the crisis of Bretton Woods System at 1968 and the oil crisis at 1974 are two structural breaks. We found 11 cointegrating vectors and 4 common stochastic trends. The price convergence of all cities is less than 5 years of half-lives when a shock occurs from its own cities, while this tends to have more than 5 years of half-lives when a shock happens from other cities. The interest rate is an effective tool for controlling cities' price in short run.**

***Index Terms*— Half-life, cointegration, U.S. cities**

I. Introduction

The price stability is one of the criteria and this led researches to focus on the price convergence in the European Union, such as proven by Camarero et al.(2000), Trivez (2001), Roger (2001), Holmes (2002), Sosvilla-Rivero and Gil-Pareja (2004), Friberg and Matha (2004), Jenkins (2004), Goldberg and Verboven (2005), Robinson (2007), Rogers (2007), Gil-Pareja and Sosvilla-Rivero (2008), and Parsley and Wei (2008). It is known that the objective of monetary policy of European Central Bank (ECB) is price stability. This is because the price stability contributes to achieving high levels of economic activity and employment. However, although researches on price convergence have been extensively studied, most of them were not concerned how the monetary policy impact to price convergence.

Rogers (2007) found that the price dispersion in the European countries is quite close to that of US cities. This supports the idea of Cecchetti et al. (2002) that the studying the behavior of prices across U.S. cities will help us in understanding the likely nature of inflation convergence in the Euro area. This implies that the studying of prices across cities in a developed country can be considered as an example of inflation convergence in an economic union of developed countries. Carrion-i-Silvestre et al. (2004) adopted this idea by examining price convergence in Spanish cities, Vargas-Tellez (2008) and Sonora (2005) studied the prices convergence in Mexican cities, and these studies can be considered as an example of prices convergence in an economic union of middle-income countries. Similar idea was also adopted to an economic union of developing countries, for example, Morshed et al. (2006) investigated the consumer price behavior of 25 major cities in India, Rangkakulnuwat and Ahn (2006) examined the consumer price index convergence of 5 regions in Thailand

The contribution of this paper is that we examine how the monetary policy affects to the price convergence in an economic union of developed countries. We hope that understanding this study would be helpful in implementing the monetary policy of the union.

Manuscript received October 28, 2008.

Rangkakulnuwat, P. is with University of the Thai Chamber of Commerce, Bangkok, 10400 Thailand (phone: 662-697-6305; fax: 662-277-4359; e-mail: poomthan_r@yahoo.com).

Morshed, AKM M. is with Southern Illinois University, Carbondale, IL, USA (e-mail: mmorshed@siu.edu).

Wang. H. H. is with Purdue University, West Lafayette, IN, USA. (e-mail: wanghong@purdue.edu).

Ahn, S. K. is with Washington State University, Pullman, WA, USA (e-mail: ahn@wsu.edu).

II. Theoretical Framework

A. Theoretical Framework

Let p_{it} is the logarithm of the price index for economy i during period t. Suppose the central bank of a union conducts monetary policy to control the price level in the union, p^*_{t+k}, to the direction they need. Hence, we propose the definition of convergence in price stability for price level in country i valued at information available at time t (I_t) as

$$\lim_{k\to\infty} E(p_{i,t+k} - p^*_{t+k}|I_t) = 0 \quad (1)$$

Unfortunately, p^*_{t+k} is not completely controlled by the central bank of a union: there are many macroeconomic factors that can affect the price level of the union. Therefore the p^*_{t+k} can be considered as the common trend driving the price level in other countries. If $p_{i,t+k} - p^*_{t+k}$ is a mean zero stationary process, then this definition of convergence to price stability will be satisfied.

Since price levels in countries 1, 2, …, N will converge to price stability, hence, they contain a single common trend. Moreover, it is possible that countries 1, 2,…, N do not converge to price stability in the sense of equation (1). In other words, they may face the same permanent shocks with different long run weights, as

$$\lim_{k\to\infty} E(\beta' p_{t+k} - p^*_{t+k}|I_t) = 0 \quad (2)$$

where β' is the cointegrating vector. There is only single central bank in an economic union, hence monetary conduct to affect the price stability p^*. From the monetary approach, the price is determined by the domestic nominal money supply and the real money demand. With real money demand depending on real income and the nominal interest rate, we may consider as

$$p^* = m^* - ky^* + hi^* \quad (3)$$

where m is the logarithm of nominal money, y is the logarithm of real income, and i is the nominal interest rate. The superscript * refers to the union. k and h are parameters. We get the relationship of price convergence and monetary policy by plugging (3) into (2) as

$$\lim_{k\to\infty} E(\beta' p_{t+k} - m^*_{t+k} + ky^*_{t+k} - hi^*_{t+k}\,|I_t) = 0 \quad (4)$$

This implies that price convergence and monetary policy are related in the sense that the series of $\beta' p_{t+k} - m^*_{t+k} + ky^*_{t+k} - hi^*_{t+k}$ is stationary. Therefore, one of statistical test of relationship between price convergence and monetary policy which is cointegration test should be tried.

III. Background and Data

The Bretton Woods system is the international monetary regime that prevailed from the end of World War II until the early 1970s. Under the Bretton Woods System, the IMF Articles of Agreement stipulated that each member country declare its par value in terms of gold. The dollar grice of gold was 35$ an ounce, and it was never changed, until the Smithsonian conference of December 1971 (Giovannini, 1988). Central Banks used gold in transactions among themselves, and intervened in the private bullion market at their own discretion. Since the private sector had no rights of official conversion of national currencies into gold (Tew, 1977). The crisis of the Bretton Woods international monetary system started from November 1967 to March 1973, where the four principal events were devaluation of sterling in November 1967, suspension of convertibility of the dollar into gold in August 1971, floating the sterling in June 1972, and second devaluation of dollar and generalized floating in February/ March 1973.

Moreover, United States also faced the oil crisis in 1973, which was the result of Organization of Petroleum Exporting Countries (OPEC) announced that they would no longer ship petroleum to nations that had supported Israel in its conflict with Egypt. OPEC-member agreed to use their leverage over the world price-setting mechanism for oil to quadruple world oil prices.

The data sets are annual started from between 1929 and 2005; Consumer Price Index (CPI) is used as the price index and obtained from the Bureau of Labor Statistics (BLS). Money supply, real income, and interest rate used in this paper are seasonally adjusted M1, real gross domestic product (GDP), and yield on corporate bond, respectively. Data on real income and interest rate are from Bureau of Economic Analysis and Board of Governors of the Federal Reserve System, respectively. Data on money supply for 1959 to 2005 are from Board of Governors of the Federal Reserve System and are spliced to Stock and Watson (1993)'s data for 1929-1958. All variables are taking natural logarithm except interest rate.

IV. Results

A. Estimating and Testing the Multiple Structural Breaks

In this paper, we used the method proposed by Bai and Perron (1998, 2003) to estimate and test the endogenous structural breaks in all series. When we adopted the approach, the maximum of 5 endogenous structural breaks, serial correlation in the errors and different variances of the residuals across segments are allowed. There are three statistics to identify the breaks; first, the sup $F_T(k)$ test, i.e. a sup F-type of the null hypothesis of no structural break versus the alternative of a fixed (arbitrary) number of breaks k; second, two maximum tests of the null hypothesis of no structural break versus the alternative of an unknown number of breaks given some upper bound, i.e. $UD_{\max}$ test, an equal weighted version; finally, the sup $F_T(l+1|l)$ test, i.e. a sequential test of the null hypothesis of l breaks versus the alternative of l+1 breaks.

The results of testing the multiple breaks of the logarithm of CPI in the 12 cities, and m^*, y^*, and i^*, are shown in Table 1. As it can be seen from the table, we found that $UD_{\max}$ tests are significant at least 10% level for all price index variables except for St. Louis. For the monetary policy variables, the $UD_{\max}$ tests are also significant at 1% level for m^* and i^*, but not significant for y^*. These imply that at least one break is present for m^* and i^* and all price indexes, except for St. Louis. Next, we consider the Sup $F_T(l+1|l)$ tests of the variables that $UD_{\max}$ tests were significant, we found that all of these are not significant for any $l \geq 1$, so we conclude that there is only one break in those variables.

The dates of a break estimated by the Bai-Perron test is 1974 for the logarithm of CPI in Atlanta, Detroit, Kansas City, Pittsburgh, Portland, and St. Louis and is 1975 for the logarithm of CPI in Chicago, Cincinnati, Los Angeles, Minneapolis, San Francisco, and Seattle. The break dates in 1974 or 1975 for the price indexes are corresponded to the oil crisis in 1973. For the monetary policy variables, the test estimate the break date is in 1971 for m^* and in 1968 for i^*. The break dates in 1971 and 1968 are linked to the end of the Bretton Woods system.

B. Unit Root and cointegration Analysis

Since we knew the break date in each variable from the previous section, it is therefore reasonable to apply the unit root test proposed by Perron (1989) when one structural break is in the series. The changing growth model (model (B) of Perron, 1989) is adopted to all one estimated break variables; while the ADF unit root test is adopted to variables with no breaks. Table 2 contains the t-statistics for the unit root hypothesis test of Perron (1989) and ADF, respectively. The results show that all variables are nonstationary.

Since all variables are nonstationary, cointegration analysis could be adopted to find long-run relationships. As we know from the previous section that one break exists in all the logarithms of CPI either 1974 or 1975, the oil crisis; while one break exists in m^* and i^* at 1971 and 1968, respectively, which corresponds to the crisis of Bretton Woods System. According to the critical values for rank tests proposed by Johansen et al. (2000) can be adopted with up to two structural breaks. We cannot simultaneously impose four breaks of 1971, 1974, 1975, and 1968 when cointegration relationship among these series is tested. Therefore, two breaks at 1974 and 1968 are considered to represent the oil crisis and the crisis of Bretton Woods system in this paper.

All series shows time trend inside, this suggests that it is suitable to use broken linear trend in cointegrating vectors, H_l, as explain in Johansen et al. (2000). The minimum value of Hannan-Quinn information criteria is adopted to select optimum lag length in the model which is three in this case. Table 3 represents the cointegration rank tests, which presents trace statistics, mean and variance of trace tests distribution used to calculate critical values by the estimated response surface from a Gamma distribution. The trace statistic shows that the null hypothesis of r=0 until r=10 are rejected at 0.05 significant levels, while the corresponding hypothesis of r= 11 cannot be rejected at 0.05 significant levels.

C. Half-Life Analysis

Using the orthogonal projection employed by Kasa (1992) to identify common stochastic trends, we decompose the p-vector X_t, which is cointegrated with rank r, into a sum of stationary (or temporary) components and common nonstationary (or permanent) components. These can be expressed as follows:

$$X_t = \beta(\beta'\beta)^{-1}\beta'X_t + (I - \beta(\beta'\beta)^{-1}\beta')X_t$$

$$= \beta(\beta'\beta)^{-1}\beta'X_t + \beta_\perp(\beta_\perp'\beta_\perp)^{-1}\beta_\perp'X_t \qquad (5)$$

where $\beta_\perp$ denotes the $p\times(p-r)$ factor loading matrix, whose number of columns $(p-r)$ is equal to the number of stochastic trends, and $\beta_\perp'\beta = 0$. The $p-r$ stochastic trends are identified by $(\beta_\perp'\beta_\perp)^{-1}\beta_\perp'X_t$.

According to Morshed et al. (2006), the half-life of the convergence to the common trend in response to the shock of its own logarithm of the price index of a city is obtained from the diagonal elements of $\beta(\beta'\beta)^{-1}\beta'\Psi_j$ and the responses to the shock of other cities are obtained from the off-diagonal elements. The half life is defined by the period in which the marginal change in the stationary component of the impulse response becomes half of the initial response.

Consider half-life for own prices shocks, we found that only 1 city, which is Portland, has the longest half-life (5 years), while, 6 cities, which are Chicago, San Francisco, Cincinnati, Pittsburg, St. Louis, and Minnesota, yield the fastest half-life (1 year). Moreover, Detroit, Los Angeles, Seattle, and Kansas City yield half-life of 2 years, and Atlanta yield half-life of 3 years. Consider half-life for own macroeconomic variables shocks, we found that the m^* and i^* have half-life of 1 year, and y^* yield half-life of 3 years.

Consider the half-life from other cities price shocks, we found that a price shock from Chicago may result in half-lives of 2 years for Seattle and of greater than 5 years for the other cities' price indexes. Moreover, it is possible that the price shock from Chicago could result in half-lives of more than 5 years for the macroeconomic variables. In most patterns of a price shock from a city appears to yield half-lives of greater than 5 years to the other cities' price indexes and the macroeconomic variables. Except for a price shock from Los Angeles, which may result in half-life of 3 years for m^*; a price shock from Atlanta, which seems to yield half-lives of 4 years for m^* and i^* and of 2 years for y^*; a price shock from Pittsburg, which contributes to half-life of 2 years for i^*; and a price shock from Kansas City appears to have half-life of 1 year for i^*.

On the other hand, a shock from macroeconomic variables may results in half-lives of less than 5 years for most cities' price indexes. For instance, a shock from i^* will result in the half-lives of 5 years for Chicago, and of 1 year for the other cities' price indexes; a shock from m^* yield half-lives of greater than 5 years for only 3 cities: San Francisco, Cincinnati, and St. Louis; a price shock from y^* could result in half-lives of greater than 5 years for San Francisco, Portland, and Minnesota. Moreover, a shock from m^* yield half-lives of 1 year for the other macroeconomic variables; a shock from y^* could result in half-live of 3 years for the other macroeconomic variables; and it is likely that a shock from i^* will result in half-live of 1 year for m^* and of 5 years for y^*.

Table 1. Bai-Perron tests of multiple endogenous structural breaks for the logarithm of CPI in the 12 cities and m^*, y^*, and i^*

	Tests statistics					
Variable	UD_{max}	Sup $F_T(2\|1)$	Sup $F_T(3\|2)$	Sup $F_T(4\|3)$	Sup $F_T(5\|4)$	Break Dates
Atlanta	10.05**	0.56	0.79	1.47	1.11	1974
Chicago	7.89*	1.99	0.88	0.08	1.61	1975
Cincinnati	12.18**	0.84	1.15	1.35	--	1975
Detroit	14.13***	1.43	0.46	1.58	--	1974
Kansas City	14.12***	0.47	0.90	0.50	--	1974
Los Angeles	8.12*	0.75	0.05	0.03	--	1975
Minneapolis	12.82***	0.53	0.35	1.07	0.86	1975
Pittsburgh	11.55**	0.67	0.84	1.01	0.30	1974
Portland	15.02***	1.21	1.81	0.23	--	1974
San Francisco	15.26***	1.73	0.61	0.54	0.14	1975
Seattle	13.14***	1.14	0.98	0.25	--	1975
St. Louis	10.62**	0.88	0.12	0.96	--	1974
m^*	33.51***	4.10	0.47	1.13	0.01	1971
y^*	6.58	1.25	0.66	0.12	--	-
i^*	34.06***	2.55	1.45	1.30	0.05	1968

Notes: ***,**, * denotes significance at 1% , 5% and 10% levels, respectively. We use Gauss codes provided by Perron's website to calculate the test statistics and critical values.

V. Conclusion

This paper aims to estimate the rate of price convergence in 12 U.S.' cities in response to a shock from a city and to a macroeconomic variable: m^*, y^*, or i^*. The rate of price convergence is explained by half-life, which is defined as the marginal change in the stationary component of the impulse response becomes half of the initial response (Morshed et. al., 2006).

We found that when a price shock happens from a city, its rate of price convergence for most of cities is less than 5 years of half-life with an average of 1.75 years; where Chicago, San Francisco, Cincinnati, Pittsburg, St. Louis, and Minnesota there exist 6 cities the fastest adjustment with half-lives of 1 year; while Portland is the slowest adjustment with half-life of 4 years. The results also imply that when a shock occurs in a particular city, the rates of convergence in other cities' price indexes and in the macroeconomic variables are persistent with at least 5 years of half-lives. Furthermore, a shock from a macroeconomic variable will result in shorter rates of price convergence for most cities; especially when a shock happens from i^*, the rates of price convergence for all cities, except for Chicago, are only 1 year of half-life.

ISBN: 978-988-17012-7-5

Table 2: Unit Root Test

Variables	Model	lag	Break	λ	t-statistics of $\hat{\rho}$
Perron (1989) test					
Atlanta	B	4	1974	0.6	-2.491
Chicago	B	5	1975	0.6	-3.296
Cincinnati	B	5	1975	0.6	-3.357
Detroit	B	7	1974	0.6	-3.292
Kansas City	B	5	1974	0.6	-3.283
Los Angeles	B	5	1975	0.6	-3.467
Minneapolis	B	5	1975	0.6	-3.592
Pittsburgh	B	5	1974	0.6	-3.180
Portland	B	7	1974	0.6	-3.670
San Francisco	B	4	1975	0.6	-2.796
Seattle	B	5	1975	0.6	-3.580
St. Louis	B	7	1974	0.6	-4.068
m^*	B	7	1971	0.6	-4.018
i^*	B	4	1968	0.5	-2.288
ADF test					
y^*		1			-3.430

Notes: Model (B) $\Delta z_t = \mu + \theta DU_t + \beta t + \gamma DT_t^* + \rho z_{t-1} + \sum_{i=1}^{k} c_i \Delta z_{t-i} + e_t$ where $DU_t = 1$,if $t > T_B$ and 0 otherwise; $D(TB)_t = 1$, if $t = T_B+1$ and 0 otherwise; $DT_t = t$, if $t > T_B$ and 0 otherwise. At $\lambda = 0.6$, the critical value of 1% and 5% significant levels are -4.57 and -3.95, respectively. At $\lambda = 0.5$, the critical values for $\rho = 0$ of Perron (1989) test at 1% and 5% significant levels are -4.56 and -3.96, respectively, and those for $\rho = 0$ of ADF are provided by E-view software.

Table 3. Cointegration Rank Test

Hypothesis	Mean	Variance	Trace Test	p-value
$r = 0$	1620.50	807.06	2594.62	0.000
$r \leq 1$	1222.99	804.93	2108.13	0.000
$r \leq 2$	952.29	787.97	1673.23	0.000
$r \leq 3$	761.12	757.01	1344.19	0.000
$r \leq 4$	621.18	713.61	1076.66	0.000
$r \leq 5$	514.98	659.92	878.11	0.000
$r \leq 6$	431.38	598.49	692.50	0.000
$r \leq 7$	363.16	532.10	558.36	0.000
$r \leq 8$	305.54	463.49	441.53	0.000
$r \leq 9$	255.45	395.25	329.76	0.000
$r \leq 10$	210.93	329.57	251.52	0.016
$r \leq 11$	170.87	268.22	183.52	0.216
$r \leq 12$	134.76	212.42	118.21	0.875

REFERENCES

[1] Bai , J. and Perron, P. (1998) Estimating and Testing Linear Models with Multiple Structural Changes, *Econometrica*, **66**, 47-78.
[2] Bai , J. and Perron, P. (2003) Computation and Analysis of Multiple Structural Change Models, *Journal of Applied Econometrics*, **18**, 1-22.
[3] Bernard, A. B. and Durlauf, S. N. (1995) Convergence in International Output, *Journal of Applied Econometrics*, **10**, 97-108.
[4] Camarero, M., Esteve, V., and Tamarit, C. (2000) Price Convergence of Peripheral European Countries on the Way to the EMU: A Time Series Approach, *Empirical Economics*, **25**, 149-168.
[5] Carrion-i-Silvestre, J. L., Barrio, T. D., and Lopez-Bazo, E. (2004) Evidence on the Purchasing Power Parity in a Panel of Cities, *Applied Economics*, **36**, 961-966.
[6] Cecchetti, S. G., Nelson, C. M,, Sonora, R. J. (2002) Price Index Convergence Among United States Cities, *International Economic Review*, **43**, 1081-1099.
[7] Evans, P. and Karras, G. (1996) Convergence Revisited, *Journal of Monetary Economics*, **37**, 249-265.
[8] Fendel, R. and Frenkel, M. (2008) Inflation differentials in the Euro area: did the ECB care? *Applied Economics (forthcoming)*
[9] Friberg, R., Matha, T. Y. (2004) Does a common currency lead to (more) price equalization? The role of psychological pricing points, *Economics Letters*, **84**, 281-287.
[10] Gil-Pareja, S. and Sosvilla-Rivero, S. (2008) Price Convergence in the European Car Market, **40**, 241-250.
[11] Giovannini, A. (1988) How Do Fixed-Exchange-Rates Regimes Work: The Evidence From the Gold Standard, Bretton Woods and the EMS, *NBER Working Paper No. 2766.*
[12] Goldberg, P. K., Verboven, F. (2005) Market Integration and Convergence to the Law of One Price: Evidence from the European Car Market, *Journal of International Economics*, **65**, 49-73.
[13] Holmes, M. J. (2002) Panel data evidence on inflation convergence in the European Union, *Applied Economics Letters*, **9**, 155-158.
[14] Jenkins, M. A. (2004) Purchasing Power Parity and the Role of Traded Goods: Evidence from EU states, *Applied Economics*, **36**, 1371-1375.
[15] Morshed, A.K.M., Ahn, S. K., Lee, M. (2006) Price Index Convergence Among Indian Cities: A Cointegration Approach, *Journal of Asian Economics*, **17**, 1030-1043.
[16] Parsley, D. and Wei, S. (2008) In Search of a Euro Effect: Big Lessons from a Big Mac Meal? *Journal of International Money and Finance*, **27**, 260-276.
[17] Perron (1989) The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis, *Econometrica*, **57**, 1361-1401.
[18] Perron, P. (1997) Further Evidence on Breaking Trend Functions in Macroeconomic Variables, *Journal of Econometrics*, **80**, 355-385.
[19] Rangkakulnuwat, P. and Ahn, S. K. (2006) Price Index Convergence in Thailand, *International Economics and Finance Journal*, **1**, 233-245.
[20] Robinson, T. (2007) The Convergence of Electricity Prices in Europe, *Applied Economics Letters*, **14**, 473-476.
[21] Roger, J. H. (2001) Price Level Convergence, Relative Prices, and Inflation in Europe, *International Finance Discussion papers*, vol. 699, Washington: Board of Governors of the Federal Reserve System.
[22] Roger, J. H. (2007) Monetary Union, Price Level Convergence, and Inflation: How Close is Europe to the USA?, *Journal of Monetary Economics*, **54**, 785-796
[23] Sonora, R. J. (2005) City CPI convergence in Mexico, *Review of Development Economics*, **9**, 359-367.
[24] Sosvilla-Rivero, S., Gil-Pareja, S. (2004) Price Convergence in the European Union, *Applied Economics Letters*, **11,** 39-47.
[25] Stock, J.H. and Watson, M. W. (1993) A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, *Econometrica*, 61, 783-820.
[26] Tew, B. (1977) The Evolution of the International Monetary System 1945-1977, New York: John Wiley & Sons.
[27] Trivez, F. J. (2001) Analysis of the long-term relationships of the underlying rates of inflation in the EMU member states, *Applied Economics,* **33**, 2001-2007.
[28] Vargas-Tellez, C. O. (2008) Purchasing Power Parity Across Mexican Cities: a Panel Data Analysis, *Applied Economics,* 99999:1, 1-9