Hysteresis in East Asian Unemployment

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Abstract

High-performing Asian economies are quite distinctive with efficient, flexible, and responsive labor markets. Comparing the persistence of unemployment in East Asian economies to that in Western countries is difficult due to the data's short time spans and possible structural breaks. This paper employs advanced unit-root tests to deal with these problems and fails to reject "hysteresis" in the unemployment rates of the high-performing Asian countries, even after taking into account structural change. An alternative explanation of different productivity growth for the hysteresis phenomenon of the Asian countries is proposed herein.

JEL classification: C12, C13, E24

Key words: Unemployment rate; Unit root; Hysteresis; Structural breaks

1. Introduction

Hysteresis, a term used by Blanchard and Summers (1986) to describe very high persistence or unit root in European unemployment, has attracted much attention from economists. Many researchers also find empirical evidence for unemployment hysteresis in several OECD countries.[1] Nevertheless, little work has been undertaken about unemployment in countries with market mechanisms that guide labor allocation such as that in high-performing East Asian countries. The World Bank (1993) defines high-performing Asian economies as those Asian economies, led by Japan, that have several common characteristics, such as very rapid export growth and very rapid economic growth. One important factor contributing to their growth is the existence of an efficient, flexible, and responsive labor market, and their rapid growth itself is a key factor in making their own wages more flexible.[2] Is the stationarity of unemployment rates in these East Asian economies distinct from those in Europe?

Export push and relatively high productivity are two common facts of the

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[1] See, for instance, Brunello (1990), Neudorfer, Pichelmann, and Wagner (1990), Mitchell (1993), Jaeger and Parkinson (1994), Røed (1996), and so on.

[2] There are no strong unions in these countries except for Japan. The governments of Singapore and Taiwan have suppressed independent unions while Hong Kong and Japan do not intervene in labor relations. However, most of Japan's unions are company-based and do not extract economic benefits for their members at the expense of economic growth.

high-performing East Asian countries. Theoretically, the relative price between domestic and foreign products is a key determinant of employment in small-open economies (see e.g. Turnovsky, 1995 and Razin and Yuen, 2002). On the other side, different productivity growth inducing the real exchange rate/terms of trade to be non-stationary not only is a theoretical hypothesis known as the Balassa-Samuelson effect, but also has its own empirical success.[3] It is quite reasonable that non-stationarity characteristics pass from relative prices to unemployment rates. If this is indeed so, then we can find "hysteresis" in the unemployment rates of competitive labor markets for the high-performing East Asian countries.

One reason that relatively few studies have been conducted on the stationarity of Asian unemployment rates is due to a lack of data. Except for Japan, very limited unemployment rate data series exist for high-performing Asian economies. It is well known that the power of the conventional unit-root test is low when the sample span is short. A panel-based procedure such as in Im, Pesaran, and Shin (2003) and Maddala and Wu (1999), on the other hand, is attractive since it allows one to improve the estimation efficiency by exploiting a cross-sectional variation of data.

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[3] Related works can be seen in Kim (1990), Ceglowski (1996), Ito, Isard, and Symansky (1999), and Lee, Nziramasanga, and Ahn (2002). However, some authors think that the effect of different productivity growths on the relative price can be modeled by time trend and thus interpret that Balassa-Samuelson effect implies the real exchange rate is stationary with a deterministic trend. See, for examples, Obstfeld (1993), Drine and Rault (2003), and Peel and Venetis (2003).

A structural break is also an obstacle in detecting the stationarity of the unemployment rate.[4] During East Asian economies' high growth process, fast structural changes have co-existed in their labor markets. Millions of workers have left the agricultural sector and moved into the manufacturing sector. Such rapid growth of productivity leads to rapid growth in real wages and extraordinary dynamics of labor demand. The recent Asian financial crisis seems to have become another predominant shock that caused a structural change in these countries. It appears to be encouraging that one is able to consider structural breaks in testing the unit-root hypothesis of Asian unemployment rates.

Previous panel tests, however, do not consider structural changes when testing a unit-root hypothesis. Im, Lee, and Tieslau (2005) point out that any small-sized distortion in an individual time series accumulates in the panel framework and extends the univariate LM unit-root test with level shifts to a panel LM test. Like the Im, Pesaran, and Shin (2003) test, the panel LM test is based on the pooled likelihood function and is calculated as the average of the individual LM statistics from each time series. Im, Lee, and Tieslau (2005) further show that the asymptotic distribution of the panel LM test with level shifts does not depend on the nuisance parameter indicating the position of the break. This paper therefore applies Im, Lee,

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[4] Perron (1989) argues that the unit-root hypothesis is hard to reject by the conventional ADF test if a series is comprised of a stationary fluctuation around a trend function, which contains a one-time break.

and Tieslau (2005)'s panel LM test to Asian unemployment rates.

The results presented herein fail to reject that the unemployment rates in high-performing Asian countries are integrated of order one even after taking into account the fact that many of them are subject to a structural change in their mean. A hypothetical model of Lee (2005) is provided as a preliminary explanation for the findings. The remainder of the paper is organized as follows. Section 2 provides a brief data description, showing that they fail to reject the unit-root hypothesis with the usual ADF statistic tests. Section 3 briefly describes the model and econometric strategy — univariate unit-root test and panel unit-root test, both with breaks. Section 4 reports the empirical results from the tests. Critical values are simulated to correct for a small sample bias. An important issue that we consider is the appropriate critical values for the panel data unit-root test in the presence of serial correlation and contemporaneous correlation. Section 5 provides a hypothetical model and some initial evidence. Finally, the last section summarizes the conclusions.

2. Data Description

The empirical period begins in 1976 and ends in 2004. The data includes annual observations of the unemployment rates in 9 high-performing Asian economies - Hong Kong, Indonesia, Japan, South Korea, Malaysia, the Philippines, Singapore, Taiwan, and Thailand. Our dataset comes mainly from the *International Labour Organization* and various government statistics publications. The unemployment rates of China, a country which is a relatively newer high-performing economy, are not included in our panel due to the short span in the data.[5] Figure 1 illustrates the unemployment rates of the countries in our panel. At first glance, not even one of the unemployment rates plotted appears to be a stationary series.

For the panel under investigation, Japan has low rate of unemployment relative to other industrialized economies until the beginning of the 1990s. Along with the declining importance of lifetime employment in Japanese companies, Japan's unemployment rate increases to approach the levels in other non-Euro OECD countries. For the other countries in the panel, there are substantial improvements in wage and labor productivity as labor continues to shift out of the agriculture and other low-income activities. Due to a rising share of female workers with lower unemployment rates (the rate of female workers is lower than that of male workers in South Korea), a falling share of young workers with higher unemployment rates, and a lower degree of labor market imperfections, the unemployment rate of South Korea

[5] Wu (2003) finds persistence in China's unemployment rate during the period of 1978-1998. However, the data series crosses the period of China being a Communist nation and China's openness to markets. The definition of unemployment rate has therefore fundamentally changed during the sample period.

shows a downward trend except for the period disrupted by the Asian financial crisis (Hahn, 1996). The other little tigers — Hong Kong, Singapore, and Taiwan — are less affected by the 1997 financial crisis. However, the integration of Hong Kong's economy and Taiwan's economy with China has changed the patterns of their production and trade. The infinite elasticity of China's labor supply has obviously contributed to increasing the unemployment rates of the two tigers in the late 1990s.

Unemployment has remained high in the Philippines, since she did not generate enough employment opportunities in higher productive non-agricultural jobs to absorb the rising tide of new labor force entrants caused by a high population growth rate. Her political instability is also responsible for her economic stagnation. The Philippines is the only Asian country where there has been both a large temporary and permanent outflow of skilled personnel. Indonesia is the other economy where unemployment is extremely high. The rising unemployment rate of Indonesia since the 1990s corresponds to her progress in labor rights. In Malaysia, on the other hand, unemployment appears to be cyclical in nature.[6]

To give a preliminary impression, this paper first applies the conventional ADF test to examine hysteresis in Asian countries. Our autoregression estimation applies

[6] Some detailed descriptions about labor markets in Hong Kong, Indonesia, South Korea, Malaysia, the Philippines, Singapore, Taiwan, and Thailand can be found in Manning and Fong (1990) and Manning (1999).

one model that has a constant along with another that has both a constant and a time trend.[7] The finite sample critical values of the ADF test are simulated by a non-parametric bootstrap. From panel A of Table 1 we find that the unit-root hypothesis of unemployment rates cannot be rejected by the ADF test at the 5% level of significance for the nine Asian countries considered. Although we usually do not think that unemployment rates have a time trend, an ADF test of a model with a trend is employed as a complement. However, the findings from panel B of Table 1 still indicate the high-performing Asian countries are non-stationary, except for Malaysia which rejects the hypothesis of the unit root with a trend.

3. Econometric Methodology

This paper applies two univariate unit-root tests and three panel unit-root tests. Apart from the traditional ADF test and the Im, Pesaran, and Shin (2003) and Maddala and Wu (1999) panel tests, this paper introduces Lee and Strazicich (2003)'s minimum LM test as an alternative univariate unit-root test and includes the Im, Lee, and Tieslau (2005) method as well to construct panel data unit-root tests.

Consider countries 1 to N. Let u_{nt} denote the unemployment rate in country n at time t with the following data-generating process:

[7] The lag order of the ADF regression model is selected based on the recursive *t*-statistic procedure suggested by Ng and Perron (1995) with an upper bound of $p_{\max} = 4$ on p_n.

$$u_{nt} = y_{nt}\Phi_n + \eta_{nt}, \quad n = 1,..., N, \tag{1}$$

where y_{nt} is a vector of exogenous variables, $A_n(L)\eta_{nt} = \zeta_{nt}$, with $A_n(L)$ being finite-order polynomials, and ζ_{nt} is an unexpected disturbance assumed to be uncorrelated over time and over countries.

In the basic case whereby $y_{nt} = [1 \quad t]$ in equation (1), suppose that there is a group of N unemployment rates in a panel which have the following time-series representation:

$$\Delta u_{nt} = \alpha_{0n} + \alpha_{1n}t + \beta_n u_{nt-1} + \sum_{j=1}^{p_n} \gamma_{nj}\Delta u_{nt-j} + \zeta_{nt}. \tag{2}$$

In this paper $\Delta x_t \equiv x_t - x_{t-1}$ for any variable x. Both the Im, Pesaran, and Shin (2003) test and the Fisher test examine the null hypothesis:

$$H_0: \ \beta_1 = \beta_2 = ... = \beta_N = 0,$$

against

$$H_a: \ \beta_n < 0 \ \text{ for some } \ n \in N.$$

Equation (2) neglects the likely impact of structural changes on the test of the unit-root hypothesis, however. The labor market structure's fast changes in the past thirty years are observed for high-performing Asian countries. To take into account a one-time structural change, Lee and Strazicich (2003) define $y_{nt} = [1 \quad t \quad d_{nt}]$, where $d_{nt} = 1$ for $t \geq T_{nB} + 1$, and is zero otherwise in equation (1). In the case of allowing for a change in level, we revise the ADF regression in equation (2) so that the LM unit root test statistic can be estimated by a regression according to the LM (score) principle as follows:

$$\Delta u_{nt} = \Delta y_{nt}\Psi_n + \beta_n \tilde{S}_{nt-1} + \sum_{j=1}^{p_n} \gamma_{nj} \Delta \tilde{S}_{nt-j} + \xi_{nt}, \tag{3}$$

where $\tilde{S}_{nt} \equiv u_{nt} - \tilde{\psi}_{nx} - y_{nt}\tilde{\Phi}_n$, $\tilde{\Phi}_n$ is the OLS coefficient vector in the regression of Δu_{nt} on Δy_{nt}, $\tilde{\psi}_{nx}$ is given by $u_{n1} - y_{n1}\tilde{\Phi}_n$, and ξ_{nt} is a residual.

This setting of the null model is contrary to the ADF-type endogenous break tests, which typically do not allow for the possibility of breaks under the unit root null hypothesis and suffer from spurious rejection (Nunes, Newbold, and Kuan, 1997). Let us define the univariate LM statistic of $\hat{\beta}_n = 0$, based on T observations, in (3) as $LM_{nT}^{B}(p_n)$. To jointly determine both the break points and the number of lagged differenced terms, we first determine the length of lag order at each combination of break points. The lag order p_n is again selected based on a recursive t-statistic procedure. To determine the break points T_{nB} endogenously in each time series, Lee and Strazicich (2003) utilize a procedure employed in the "minimum LM test." This means one can use a grid search to determine the breaks at the location where the t test statistic is minimized.

Lee and Strazicich (2003) prove that critical values in the minimum LM unit root test for the crash model are invariant to both the magnitude and the location of the breaks. Moreover, whether or not there are breaks, the asymptotic distribution of the LM statistic is unaffected, thereby making the minimum LM test robust to a possible mis-specification of the break point under the null. However, the small sample distribution of the LM statistic is unknown. This article simulates the critical values of these statistics to take into account the possible breaking points. The optimal number of breaks in each country is determined by sequentially examining the *t*-statistic for each break coefficient to see if it is significant at the approximate 5% level in an asymptotic normal distribution.[8] The corresponding LM unit root test statistic is then chosen after determining the optimal number of breaks. Appendix 1 gives a detailed description of our simulation procedure.

Im, Lee, and Tieslau (2005)'s panel LM test examines the null hypothesis:

$$H_0: \ \beta_1 = \beta_2 = \ldots = \beta_N = 0,$$

with the presence of a structural break, against

$$H_a: \ \beta_n < 0 \ \text{ at least for one } \ n \in N.$$

Following Im, Lee, and Tieslau (2005), we denote the panel LM test statistics as:

$$\Gamma_{LM}^{B}(p) = \frac{\sqrt{N}\left\{\overline{LM}_{NT}^{B}(p) - \sum_{n=1}^{N} \mathrm{E}[LM_T(p_n)]/N\right\}}{\sqrt{\sum_{n=1}^{N} \mathrm{Var}[LM_T(p_n)]/N}}, \qquad (4)$$

where $\overline{LM}_{NT}^{B}(p) = \sum_{n=1}^{N} LM_{nT}^{B}(p_n)/N$, and $\mathrm{E}[LM_T(p_n)]$ and $\mathrm{Var}[LM_T(p_n)]$ are

[8] We begin with the two-break LM test. If both breaks are not significant at the same time, we then employ the one-break LM unit root test for the investigated unemployment rate. If no break is significant in the one-break LM unit root test, then we further employ the no-break LM unit root test of Schmidt and Phillips (1992).

the mean and variance of $LM_{nT}^{B}(p_n)$, respectively, which are listed for various T and p in Im, Lee, and Tieslau (2005)'s Table 1. The limiting distribution of the standardized LM test statistic with or without breaks in equation (4) is a standard normal.

O'Connell (1998) shows that the presence of cross-sectional correlation in the panel data unit root test context causes size distortions. Im, Pesaran, and Shin (2003) suggest demeaning across the cross-sectional units can solve the problem, yet Jönsson (2006) argues that demeaning causes size distortion for general forms of cross-sectional correlation.[9] Another reason for this article not using demeaned data in our panel data unit-root tests is that the breaking point in one series in general is different to another series' breaking point and demeaning will pass a break from one series to the rest of the series in the panel. To take into account that residuals in equation (3) might be contemporaneously and serially correlated and to improve the small-sample performance of panel unit-root tests, this paper simulates the small-sample distribution of the Γ_{LM}^{B} statistic using the block-sampling bootstrap method. Appendix 2 offers a detailed description of the simulation procedure.

4. Empirical Investigation

4.1 Univariate Unit-root Tests

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[9] Other discussions about cross correlation in panel data unit-root tests can be referred to Chang (2004), Smith et al. (2004), and Pesaran (2007).

Since the ADF test does not allow for possible structural breaks, the failure of rejecting the unit root hypothesis may be due to a structural change, which is an important characteristic of the labor markets of high-performing Asian economies. As a consequence, we re-examine the unit-root hypothesis with Lee and Strazicich (2003)'s minimum LM test with breaks in mean.

The LM unit root tests are comparable to the corresponding Dickey-Fuller type endogenous break tests of Zivot and Andrews (1992), but are not subject to spurious rejections under the null. In each test the break points are determined endogenously from the data via a grid search by selecting the break where the unit root test statistic is a minimum (i.e., the most negative). Using the minimum LM tests of Lee and Strazicich (2003), the unit root test statistic is estimated at each combination of break points. The procedure is repeated over the time interval $[.1T, .9T]$ to eliminate end points, until the breaks are determined where the unit root *t*-test statistic is minimized. The finite sample distribution values of the univariate LM test depending on the optimal lag length (p_n) and breaking point (T_{nB}), if significant, are constructed using the bootstrap.

When employing the two breakpoints of the LM unit root test of Lee and Strazicich (2003), no country under consideration is found to have two significant breaks at a 5% level of significance, but some of them have one break and some have none. The corresponding LM unit root test statistic is then computed after determining whether a break exists or not. If no break is significant at a 5% level, then we employ the no-break LM unit root test of Schmidt and Phillips (1992); otherwise, Lee and Strazicich (2003)'s method is employed. The corresponding LM unit root test statistic is then computed after determining whether a break exists or not.

After taking the likely structural changes into account, findings from Table 2 still indicate the failure of rejecting the unit-root hypothesis at a 5% level of significance, except for Malaysia's and the Philippines' unemployment rates. The finding about Malaysia's unemployment is consistent with that in Table 1, which indicates that the unemployment rate of Malaysia is trend stationary. The finding of the unit-root test about the unemployment rate of the Philippines cannot reject its stationarity with a break. The estimated break in 1983 for the Philippines appears to reflect changes in the industrial structure and economic reforms of the country. It happens to be the year of the assassination of Senator Aquino, which turned out to be a direct cause of a severe political and economic crisis in the Philippines. The finding also indicates the lower power of unit-root tests that do not consider structural changes.[10]

[10] Likewise, Baddeley, Martin, and Tyler (1998) suggest that the apparent stochastic non-stationarity in British unemployment over 1965-1995 is in fact due to an upward structural shift in the mean rate, and

The unemployment rate in Hong Kong is non-stationary, even though there is a significant breaking point of it in 1997 corresponding to the slowing down of Hong Kong's economic growth after being handed back to mainland China as well as being attacked by the Asian financial crisis. There is no statistically significant break identified in the unemployment rate series of Japan, Indonesia, South Korea, Malaysia, Singapore, Thailand, and Taiwan. However, note that when the magnitude of the break is small, the break point cannot be correctly estimated, but the minimum LM test does not actually suffer from a significant loss of power in this case.

4.2 *Panel Data Unit-Root Tests*

The failure of rejecting the unit-root hypothesis of unemployment rates may be due to the tests' low power in the short-span sample. We therefore apply the Im, Perasan, and Shin (2003) test and the Fisher test to re-examine the stationarity of Asian unemployment rates. Maddala and Wu (1999) also find that the Im, Perasan, and Shin (2003) test suffers from considerable size distortions when residuals are cross-correlated. To take into account the effect of a contemporaneous correlation of data on critical values, we simulate the finite sample critical values of the test with a non-parametric bootstrap through 100,000 iterations.

Without taking structural changes into account, findings from panels A and B of

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Kratena (2000) concludes that the persistence of Austrian unemployment in 1964-1996 may arise from a stepwise stationary unemployment rate.

Table 3 indicate that the without trend unit-root hypothesis of Asian unemployment rates fails to be rejected by the Im, Pesaran, and Shin (2003) test and the Fisher test, respectively, at the 5% level of significance with block sizes 4, 6, and 8. This finding of stationarity is consistent with that in Christopoulos**,** Loizides, and Tsionas (2005)'s study of European countries, but is contrary to that in Song and Wu (1997, 1998)'s research on American states and OECD countries as well as Camarero and Tamarit (2004)'s research on the OECD. Although the Im, Pesaran, and Shin (2003) test and the Fisher test have a higher power relative to the conventional single-equation based test, it fails to take structural changes into account. Therefore, it is reasonable to doubt that the failure of rejecting the unit-root hypothesis, in spite of increased power from the panel data by the Im, Pesaran, and Shin (2003) test and Fisher test, may be due to its failure to incorporate likely structural changes.

Considering a one-time structural change in panel data unit-root tests, we construct panel data unit-root statistics, $\Gamma_{LM}^{B}(p)$, with heterogeneous breaking points in the panel. To allow for a contemporaneous and serial correlation in residuals, we construct the critical values of the panel LM test statistics by a block-sampling bootstrap. Findings from panel A of Table 4 point out that the unit-root hypothesis of unemployment rates is rejected at the 5% level with block sizes of 4, 6, and 8. This finding seems to reflect the higher power of the panel unit test of minimum LM statistics when break points exist.

Karlsson and Lothgren (2000) point out that a rejection of the joint unit-root null can be driven by a few stationary series. To detect whether the whole panel is stationary, we exclude those countries with stationary unemployment rates from our panel in Table 2. After excluding the Philippines and Malaysia, findings from panel B of Table 4 show that the unit-root hypothesis of unemployment rates cannot be rejected at the 5% significance level invariant to block sizes of 4, 6, or 8 now. If we further take Japan and Thailand away from the panel, findings from panel C of Table 4 provide stronger evidence for the hysteresis of the unemployment rates in Asia in terms of the p-values of the panel unit-root test statistics.

In summary, after taking into account the fact that some of unemployment rates of high-performing Asian countries are subject to a structural change in mean, we fail to provide evidence to reject that these rates are integrated of order one at the usual significance level, except for Malaysia and the Philippines. Our finding is an interesting contrast to those in previous empirical studies claiming that many Western unemployment rates are indeed stationary, but have structural changes (see e.g. Bianchi and Zoega, 1997, 1998; Lumsdaine and Papell, 1997; Katsimi, 2000; Papell, Murray and Ghiblawi, 2000; and Arestis and Biefang-Frisancho Mariscal, 2000).

5. A Hypothetical Model for I(1) Unemployment Rates

One possible model that explains the non-stationarity of unemployment rates is in Lee (2005), which extends Turnovsky (1995). Assuming uneven technology progress between a small-open economy and the rest of the world, she argues that an economy's terms of trade would embed a stochastic trend, and its natural rate of unemployment, affected by the terms of trade, thus has a unit root. Following Turnovsky (1995)'s labor market equilibrium model in which information is imperfect and assuming Okun's law holds, a Phillips curve of an open economy can be derived as:

$$u_t = \alpha_0 + \alpha_1(\pi_t - \pi_t^e) + \alpha_2(\mathrm{TOT}_t^e) + s_t + \mathrm{X}_t^u, \tag{5}$$

where u_t denotes the unemployment rate; α_0 reflects institution factors that may affect the unemployment; π_t denotes the inflation rate of domestic products while π_t^e is its expectation based on *t*-1 information set; TOT_t^e is the expectation of period *t* terms of trade based on *t*-1 information set; s_t denotes a negative supply shock and is measured by the long-run (potential) growth rate of the economy less its actual rate; and X_t^u is a deterministic catchall variable that stands for all other variables, including structural changes, and affects the unemployment rates.

According to basic macroeconomics, it is certain that $\alpha_1 < 0$. However, an improvement in the terms of trade generates both a substitute and an income effect, and so the sign of α_2 is unambiguous. Therefore, a co-integration relation between the terms of trade and the unemployment rate is a sufficient but not a necessary condition for the theory to hold.

After assuming the evolution process of the inflation rate and the terms of trade, Lee (2005) uses the structural model of a small-open economy to analyze the hysteresis of Taiwan's unemployment rates between 1951 and 2001. With some auxiliary assumptions on the autogressive process of exogenous shocks, the model is estimated by using the Kalman filter and tested with the maximum likelihood ratio test.

Because estimating a completely-specified structural model with a lot of free parameters usually has a degree of freedom problem and we do not have a long enough time span of data for each country in our panel, only a simple ADF unit root test for the terms of trade is performed. The terms of trade are constructed as the export price index divided by the import price index. The unit value of exports and imports data are from the *International Financial Statistics* (IFS), except that Taiwan's data are from its government website. The sample period is basically between 1976 and 2004 with one exception — Singapore. Singapore did not provide its export price data before 1979 and so its ADF t statistic is estimated in 1979-2004. In addition, among the nine Asian economies, Malaysia has ceased to report its import price data since 1988, the Philippines did not release its export and import price data before 1990, while Indonesia has never provided its import prices. The three countries are thus excluded from the estimation.

The estimating results are listed in Table A of Appendix 3. It is found that Japan, South Korea, Singapore, Taiwan, and Thailand, but not Hong Kong, have non-stationary terms of trade. Non-stationary terms of trade are the basis of the whole story. Therefore, the hypothetical model may be one of the candidates for the cause of non-stationary unemployment rates in high-performing Asian countries and it is obviously worth further investigation.

6. Conclusions

The labor markets of high-performing Asian economies are characterized with high flexibility and efficiency. Since there is a very short span of Asian data available for empirical research on unemployment hysteresis, econometric advances have been made to increase the power of unit root tests. Three distinct directions of testing have been employed: panel unit root tests, univariate unit root tests with structural break, and a panel unit-root test with structural breaks. This paper investigates whether the failure to reject the unit-root hypothesis of Asian unemployment rates is due to the low power of conventional tests or whether the unit-root hypothesis is the correct hypothesis. Except for the Philippines and Malaysia, we fail to provide evidence to reject that unemployment rates of high-performing Asian countries are integrated of order one even after taking into account the fact that some of them are subject to a structural change in mean.

Contrary to the rigid labor markets in Europe, high-performing East Asian economies are quite distinct with their very flexible and efficient labor markets. Our finding of hysteresis in unemployment rates of East Asian economies is therefore interesting and surprising. However, the preliminary finding appears to be consistent with the implication of the labor market in a small-open economy characterized with relatively high (low) productivity growth. More strict investigations such as tests for a simultaneous model or for cointegration between the unemployment rate and terms of trade are obviously worthy of further research.

Appendix 1

We provide herein a detailed description of the simulation procedure for deriving the small-sample distribution of the minimum LM statistic.

Step 1: We generate the following unemployment rate process (u_{nt}^*) for a country without a significant break in its unemployment rate:

$$u_{nt}^* = \eta_{nt},$$

where η_{nt} is generated from a driftless random walk model. For countries with breaks in their unemployment rate series that are significant, we add to the generating process an additional exogenous variable $\hat{\phi}_n d_{nt}$, with $d_{nt} = 1$ for $t \geq T_{nB} + 1$; and zero otherwise, and T_{nB}'s and $\hat{\phi}_n$'s are the estimates based on Lee and Strazicich (2003)'s LM principle estimation - that is:

$$u_{nt}^* = \hat{\phi}_n d_{nt} + \eta_{nt}.$$

Step 2: We estimate the following regression:

$$\Delta u_{nt}^* = \Delta y_{nt} \Psi_n + \beta_n^* \tilde{S}_{nt-1} + \sum_{j=1}^{p_n^*} \gamma_{nj}^* \Delta \tilde{S}_{nt-j} + \text{residual},$$

where $\tilde{S}_{nt} \equiv u_{nt}^* - \tilde{\psi}_{nx} - y_{nt} \tilde{\Phi}_n$, $\tilde{\Phi}_n$ is the OLS coefficient vector in the regression of Δu_{nt}^* on Δy_{nt}, and $\tilde{\psi}_{nx}$ is given by $u_{n1}^* - y_{n1} \tilde{\Phi}_n$. For a country without a significant break in its unemployment rate regression, $y_{nt} = [1 \quad t]$, and we compute Schmidt and Phillips (1992)'s no-break LM statistic to test $\beta_n^* = 0$. For a country with one significant break (two breaks) in its unemployment rate, $y_{nt} = [1 \quad t \quad d_{nt}]$ ($y_{nt} = [1 \quad t \quad d_{n1t} \quad d_{n2t}]$), and we compute Lee and Strazicich (2003)'s minimum LM statistic for testing $\beta_n^* = 0$ under each breaking point (combination of two breaking points) lying between 0.1*T* and 0.9*T*, and we construct the LM_{nT}^B statistic by finding the breaking point that minimizes the *t* statistics.

Step 3: Repeat the above steps 10,000 times to get the distribution of LM_{nT}^B.

Appendix 2

This appendix describes how to compute the distribution of the Γ_{LM}^{B} statistics of at most one structural break as follows.

Step 1: We estimate the following regression for a country with a significant break in its unemployment rates:

$$\Delta u_{nt} = \Delta y_{nt}\Psi_n + \beta_n \tilde{S}_{nt-1} + \sum_{j=1}^{p_n} \gamma_{nj}\Delta\tilde{S}_{nt-j} + \text{residual},$$

where $\tilde{S}_{nt} \equiv u_{nt} - \tilde{\psi}_{nx} - y_{nt}\tilde{\Phi}_n$, $\tilde{\Phi}_n$ is the OLS coefficient vector in the regression of Δu_{nt} on Δy_{nt}, and $\tilde{\psi}_{nx}$ is given by $u_{n1} - y_{n1}\tilde{\Phi}_n$. We then can achieve the no break part of the unemployment rates, η_{nt}:

$$\eta_{nt} = u_{nt} - \hat{\phi}_n d_{nt}\text{, if the break is significant,}$$

$$= u_{nt}\text{, otherwise,}$$

where $d_{nt} = 1$ for $t \geq \hat{T}_{nB} + 1$; zero, otherwise; and $\hat{T}_{nB}$'s and $\hat{\phi}_n$'s are estimated in computing the minimum LM statistics.

Step 2: We obtain the bootstrap sample of the error term $\varepsilon_t^0 = [\varepsilon_{1t}^0, \varepsilon_{2t}^0, \cdots, \varepsilon_{Nt}^0]$ by estimating the following system equations using the iterative seemingly unrelated regression (SUR) method:

$$\Delta\eta_{nt} = \alpha_n + \beta_n \eta_{nt-1} + \sum_{j=1}^{p_n} \gamma_{nj}\Delta\eta_{nt-j} + \varepsilon_{nt}^0, \quad n = 1, \ldots, N.$$

Step 3: Apply the block re-sampling procedure as described in Berkowitz and Kilian (1996) to generate residuals for simulation - that is, we divide $\varepsilon^0 = [\varepsilon_1^0, \ldots, \varepsilon_T^0]'$ into $T-k$ overlapping blocks with length $k+1$ and randomly select a block with replacement, where $\varepsilon_j^0 = [\varepsilon_{1j}^0, \ldots, \varepsilon_{Nj}^0]$, $j = 1, \ldots, T$. We first draw a pseudo-random number independently and uniformly from the interval of (0, 1) and then use it to generate a random number integer ν that takes on the value $j = 1, \ldots, T-k$ with equal probability. Once ν is generated, we draw a block of fitted residuals $\bar{\varepsilon}_1 = [\varepsilon_\nu^0, \ldots, \varepsilon_{\nu+k}^0]'$ to obtain $\bar{\varepsilon}_1^*$. Repeating this operation $m = T/(k+1)$ times yields a complete bootstrap sample of the error terms, $\varepsilon^* = [\bar{\varepsilon}_1^*, \ldots, \bar{\varepsilon}_m^*]'$. The bootstrap sample η_{nt}^* for η_{nt} is generated as:

$$\eta_{nt}^* = \eta_{nt-1}^* + \sum_{j=1}^{p_n} \hat{\gamma}_{nj} \Delta \eta_{nt-j}^* + \varepsilon_{nt}^*, \quad n = 1, \ldots, N,$$

where $\hat{\gamma}_{nj}$'s are the estimates obtained from step 2. The initial values of η_{n0}^* are obtained by block re-sampling - that is, we divide η_{nt} into $T-k$ overlapping blocks and randomly select a block with a replacement for η_{n0}^*. We then construct u_{nt}^* from η_{nt}^*:

$$u_{nt}^* = \eta_{nt}^* + \hat{\phi}_n d_{nt}, \text{ if the break is significant,}$$

$$= \eta_{nt}^*, \text{ otherwise,}$$

where d_{nt}'s and $\hat{\phi}_n$'s are obtained in step 1.

Step 4: Estimate the following regression:

$$\Delta u_{nt}^* = \Delta y_{nt} \Psi_n^* + \beta_n^* \tilde{S}_{nt-1}^* + \sum_{j=1}^{p_n^*} \gamma_{nj}^* \Delta \tilde{S}_{nt-j}^* + \text{residual}, \quad n = 1, \ldots, N,$$

where $\tilde{S}_{nt}^* \equiv u_{nt}^* - \tilde{\psi}_{nx}^* - y_{nt} \tilde{\Phi}_n^*$, $\tilde{\Phi}_n^*$ is the OLS coefficient vector in the regression of Δu_{nt}^* on Δy_{nt}, and $\tilde{\psi}_{nx}^*$ is given by $u_{n1}^* - y_{n1}\tilde{\Phi}_n^*$. For a country without a significant break in its unemployment rate regression, $y_{nt} = [1 \quad t]$, and we compute Schmidt and Phillips (1992)'s no-break LM statistic to test $\beta_n^* = 0$. For a country with a significant break in its unemployment rate, $y_{nt} = [1 \quad t \quad d_{nt}]$, we compute Lee and Strazicich (2003)'s minimum LM statistic for testing $\beta_n^* = 0$ under each breaking point lying between 0.1*T* and 0.9*T*, and we construct the LM_{nT}^B statistic by finding the breaking point that minimizes the *t* statistics.

Step 5: Construct the panel Γ_{LM}^B statistic based on the individual LM statistic from step 4.

Step 6: Repeat steps 3 to 5 a total of 5,000 times and the collection of realized Γ_{LM}^B statistics forms the bootstrap distribution of Γ_{LM}^B under the null hypothesis.

Appendix 3

Table A: ADF Unit-root Tests for the Terms of Trade

Country	*t* statistic	P-value
Hong Kong	-3.741 *	0.009
Japan	-1.795	0.375
South Korea	0.246	0.971
Singapore	0.835	0.993
Taiwan	-2.028	0.274
Thailand	-1.639	0.449

Note: The regression has an intercept, but without trend; * means significance at the 5% level.

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Table 1: ADF Unit-root Tests (1976-2004)

A. Considering intercept			
Country	Optimal lag	*t* statistic	P-value
Hong Kong	3	-0.613	0.813
Indonesia	3	3.447	1.000
Japan	3	-2.117	0.227
South Korea	0	-2.460	0.134
Malaysia	0	-1.135	0.685
Philippines	0	-0.848	0.789
Singapore	0	-2.456	0.135
Taiwan	1	-1.311	0.612
Thailand	0	-2.923 [a]	0.055
B. Considering intercept and time trend			
Country	Optimal lag	*t* statistic	P-value
Hong Kong	3	-0.679	0.929
Indonesia	3	1.643	0.999
Japan	3	-2.828	0.187
South Korea	0	-2.410	0.366
Malaysia	3	-3.736 *	0.036
Philippines	0	-3.129	0.120
Singapore	0	-2.310	0.414
Taiwan	1	-2.158	0.497
Thailand	0	-2.819	0.203

Note: * and [a] mean significance at 5% level and 10% level, respectively.

Table 2: Lee and Strazicich (2003)'s Minimum LM Test (1976-2004)

Country	Optimal lag	LM statistic	P-value	Break (T_{nB})
Hong Kong	0	-1.998	0.430	1997*
Indonesia	4	-0.800	0.941	1982
Japan	3	-2.832[a]	0.080	1993
South Korea	0	-2.455	0.201	1996
Malaysia	3	-3.137*	0.044	1989
Philippines	0	-3.430*	0.029	1983*
Singapore	0	-2.456	0.202	1987
Taiwan	1	-2.279	0.280	1995
Thailand	0	-2.834[a]	0.098	1991

Note: * and [a] mean significance at 5% level and 10% level, respectively.

Table 3: Panel Unit-root Tests without Breaks (1976-2004)

A. Im, Perasan, and Shin (2003) test		
W_{tbar} statistic	Block size	P-value
	4	0.482
-0.545	6	0.380
	8	0.435
B. Maddala and Wu (1999)'s Fisher test		
λ statistic	Block size	P-value
	4	0.209
27.855	6	0.159
	8	0.201

Note: Im, Pesaran, and Shin (2003)'s standardized *t*-bar statistic (W_{tbar}) statistic is defined as follows:

$$W_{tbar} = \frac{\sqrt{N}\left\{ \bar{t}_{NT}(p) - \sum_{n=1}^{N} \mathrm{E}[t_T(p_n)/N] \right\}}{\sqrt{\sum_{n=1}^{N} \mathrm{Var}[t_T(p_n)]/N}},$$

where $\bar{t}_{NT} = \sum_{n=1}^{N} t_{nT}(p_n)/N$, and $\mathrm{E}[t_T(p_n)]$ and $\mathrm{Var}[t_T(p_n)]$ are respectively the mean and variance of $t_{nT}(p_n)$, which is the t statistic of the ADF test for the *n*-th country. Maddala and Wu (1999)'s Fisher test statistic λ is defined as:

$$\lambda = -2\sum_{n=1}^{N} \ln PV_n ,$$

where PV_n is the p-value of the ADF test for the *n*-th country. The P-values for each test are computed from the corresponding finite sample distribution with a bootstrap through 5,000 iterations.

Table 4: Im, Lee, and Tieslau (2005)'s Panel LM Tests (1976-2004)

A. Whole panel		
Γ_{LM}^{B} statistic	Block size	P-value
	4	0.003
-2.489	6	0.004
	8	0.016
B. Panel excluding Malaysia and the Philippines		
Γ_{LM}^{B} statistic	Block size	P-value
	4	0.071
-1.228	6	0.051
	8	0.074
C. Panel excluding Japan, Malaysia, the Philippines, and Thailand		
Γ_{LM}^{B} statistic	Block size	P-value
	4	0.221
-0.201	6	0.162
	8	0.190

Note: The whole panel includes Japan, Hong Kong, Indonesia, South Korea, the Philippines, Singapore, Thailand, Malaysia, and Taiwan.

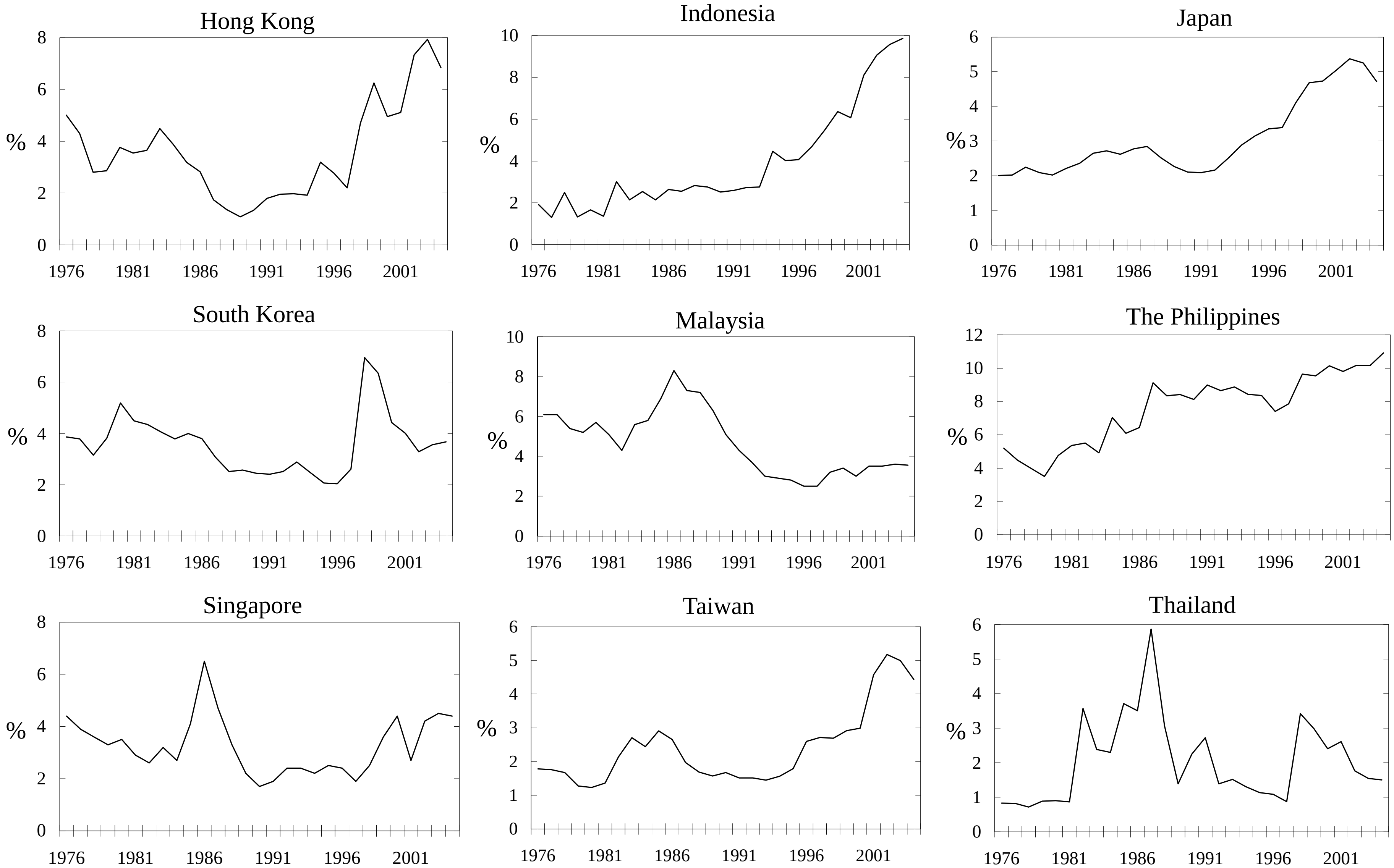

Figure 1. The Unemployment Rates in Nine Asian Economies (1976-2004)