

## Modeling China Inflation Persistence\*

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This paper employs the recently developed structural stability tests with unknown break point and two median unbiased estimation methods to model China inflation persistence over the most recent quarter of a century. Our empirical results suggest that the persistence of both consumer price inflation and retail price inflation has witnessed declines over the most recent period of low inflation, with the reduction in consumer price inflation being more pronounced. The reduced inflation persistence helps to explain the remarkable phenomenon of high economic growth under the low inflation environment in China during the recent decade.

*Key Words:* Inflation; Inflation persistence; Structural break; Monetary policy.  
*JEL Classification Numbers:* E31, E52, E58, C22.

### 1. INTRODUCTION

The People's Bank of China (PBOC), the central bank of China, specifies "stable prices" as one of its fundamental objectives, emphasizing the importance of inflation control in its monetary policy design. Not surprisingly, therefore, understanding the determinants of inflation in its dynamic process is of critical importance for the conduct of monetary policy. Modern monetary economic theories as in, for example, Clarida et al. (1999) and Stock and Watson (2002), point to inflation being determined by the

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output gap and interest rate within a dynamic system that involves an Euler equation, a short-run Phillips curve equation, and a central bank loss function. In the context of such a system, the persistence of inflation has been a regular area of study since such inertia appears to be the stylized facts observed in empirical studies (Fuhrer, 1995).

In light of this stylized facts, little research has been conducted to investigate the nature of inflation persistence in China. Nonetheless, the dynamic performance of inflation in China since the beginning of the 1980s has been particularly intriguing. In particular, the level of inflation after the end of the 1990s has been much lower than in the preceding decade. This marked changes in China's inflation experience, in particular the remarkable shift from a high and volatile inflation regime before the middle of the 1990s to low and stable inflation over the recent decade, has spurred researchers and policy makers to seek a good understanding of the inflation dynamics in China, notably Guerineau and Guillamont (2005), Gerlach and Peng (2006), and Austin et al. (2007), among many others. A general consensus achieved in these studies is that the advent of low inflation in China over the recent decade has coincided with structural changes in macro and monetary policies.

The finding in these studies raises the important question of whether the dynamic process of inflation in China has also changed in a manner which now helps to entrench low inflation. At issue, among other things, is the nature of inflation persistence. Inflation persistence measures the tendency for inflation to deviate from its equilibrium level for an extended period when perturbed (Fuhrer, 1995). Hence, if the dynamic process of inflation becomes less persistent, a given shock to the price level which boots inflation now will have a less protracted impact on the rate of ongoing inflation. Therefore, low inflation persistence entails less effort from monetary authority while high inflation persistence makes the task of reducing inflation more difficult if inflation prevails in the economy.

In view of this important implication of inflation persistence on monetary policy analysis, a small but growing literature has flourished examining the nature of inflation persistence, with the specific focus on industrial economies, notably Taylor (2000), Cogley and Sargent (2001), Cecchetti and Debelle (2006), Pivetta and Reis (2007), and Zhang, Osborn, and Kim (2008) studying the issue for the U.S., O'Reilly and Whelan (2005) for Europe, and Levin and Piger (2004) for twelve industrial countries. Surprisingly little research however, has been conducted on the study of inflation persistence in China.

The purpose of the present paper is to characterize the nature of possible structural changes in China inflation persistence over time. To achieve this objective, we employ the tests of Andrews (1993) and Andrews and Ploberger (1994) that are explicitly designed to detect structural changes

at unknown points of time. In addition, we utilize Hansen's (1999) grid bootstrap median unbiased (MU) estimator to estimate the parameter of inflation persistence. Hansen's method also provides a confidence interval with desirable asymptotic coverage. To provide a direct comparison, we also employ the MU estimator of Andrews and Chen (1994), using conventional Monte Carlo simulations, to estimate inflation persistence based on consumer price index (CPI) and retail price index (RPI) in China.

To preview our empirical results, we find that the persistence of both the CPI and RPI inflation has experienced some decline since the end of the 1990s. However, the reduction in the persistence of the CPI inflation appears to be much more pronounced in comparison to that of the RPI inflation. This finding also raises an interesting problem for future research and for the monetary authorities in China: which price index should be the primary focus when implementing a monetary policy. It is also worth noting that while most of our empirics follow the existing literature in focusing on univariate regressions, the baseline finding concerning the persistence parameter is also shown to be robust to an augmented model that takes into account the feedback from real economic growth to inflation.

This paper is organized as follows. Section 2 utilizes two median unbiased estimators to analyze the statistical nature of inflation persistence over an entire sample of 1981-2007 and provides rolling estimates of the persistence parameter in a univariate autoregressive (AR) model. Section 3 provides formal structural break tests for the dynamic model of inflation and presents our empirical results. In section 4, we examine the sensitivity of the baseline finding by taking into account possible feedback from the real economic growth to inflation. Section 5 discusses policy implications of the baseline finding and section 6 concludes the paper.

## 2. FULL SAMPLE AND ROLLING ESTIMATIONS

### 2.1. The Data

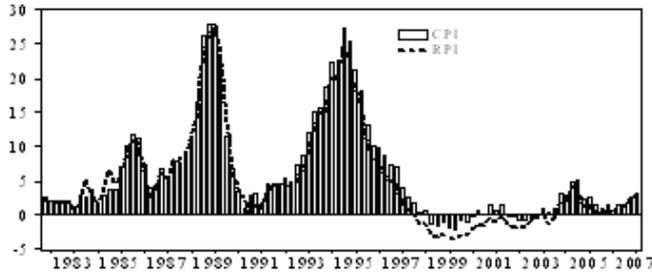
The data for the inflation series considered in this paper are chosen to provide relations that are of most interest for monetary policy analysis and to facilitate comparisons with the relevant literature. In all, we consider quarterly data on two commonly used inflation variables in China over 1981Q1-2007Q1, namely the year-on-year growth rate of the consumer price index denoted CPI and secondly the retail price index, RPI. The data are from National Bureau of Statistics of China and the sample period is dictated by data availability.

These two inflation measures are widely considered important in the study of the macro economy of China. The monetary policy committee in China monitors closely the CPI, and also monitors the information contained in the RPI, which notably excludes price indices of services. For the

Chinese economy, the use of the RPI is attractive since it limits the potential impact of price regulations on inflation developments and provides a means of comparing the nature of inflation persistence between different inflation measures.

Figure 1 displays the two inflation series by quarters from 1981 to 2007. The figure suggests that inflation in China was relatively low in the early 1980s, but trended upwards over the periods of the late 1980s and the middle 1990s, then dropped sharply by the late 1990s and remains low and stable since then. Further inspection reveals that the CPI and the RPI do not behave exactly the same across time. In particular, the RPI appears to be persistently lower than the CPI since the end of the 1980s and the difference is more pronounced during 1996-2001. The figure also shows that the CPI is noisier than the RPI, especially after the end of the 1990s.

FIG. 1. China CPI and RPI inflation series: 1981Q1-2007Q1



Data source: National Bureau of Statistics of China.

## 2.2. The Model and Estimation Methods

The model we use to study the statistical nature of inflation persistence follows the customary AR process in the literature (e.g., Andrews and Chen, 1994), namely:

$$\pi_t = c + \rho\pi_{t-1} + \sum_{i=1}^{p-1} \alpha_{\Delta i} \Delta \pi_{t-i} + u_t, \quad (1)$$

where  $c$  denotes a constant,  $\pi_t$  is the rate of inflation,  $\Delta \pi_{t-i} = \pi_{t-i+1} - \pi_{t-i}$ ,  $p$  is the optimal lag order specified for the AR model, and  $u_t$  is a serially uncorrelated error term.

Notice that the representation in model (1) summarizes the impact of past levels of inflation on current inflation through the single coefficient  $\rho$ . By construction, the parameter  $\rho$  measures inflation persistence and can be estimated with sufficient precision through least squares, even if the

individual coefficients in the dynamic process are less precisely estimated due to possible collinearity between the lagged values of inflation; provided the true data generating process under equation (1) does not contain a unit root and the sample size is sufficiently large.

However, it is well known (Phillips, 1977) that the Ordinary Least Squares (OLS) estimate of  $\rho$  tends to be biased downwards towards zero particularly when the true value of the persistence parameter is close to unity. In addition, the conventional asymptotic confidence interval associated with the estimate is also imprecise. To compensate for this problem, Andrews and Chen (1994) and Hansen (1999), among others, propose bias corrections for the estimate when it approaches one.

The Andrews-Chen estimator attempts to find the approximately median-unbiased estimator (and the associated confidence interval) for the autoregressive parameter in AR models. This MU estimator is generated from a transformation of an initial estimator, and corrected for the downward bias in conventional (least-squares) estimation through Monte Carlo simulation. A notable assumption of the Andrews-Chen method is that the disturbance term in the model is assumed to be i.i.d. Gaussian.

Hansen's (1999) grid bootstrap method extends the contribution of Andrews and Chen (1994). It calculates the bootstrap distribution over a grid of values of the autoregressive parameter in AR models and then forms the confidence interval by following a no-rejection principle. The  $\beta$ -level confidence interval can be written as

$$C = \{\rho \in R : q_T^*(\theta_1|\rho) \leq S_T(\rho) \leq q_T^*(\theta_2|\rho)\} \quad (2)$$

where  $q_T^*(\theta_1|\rho)$  denotes the bootstrap quantile function,  $S_T(\rho)$  is a non-degenerate test statistic of the null hypothesis of  $\rho$ ; being a specified value,  $\theta_1 = 1 - 0.5(1 - \beta)$ , and  $\theta_2 = 0.5(1 - \beta)$ . We use the associated 50 percent percentile estimate to derive the grid bootstrap MU estimate.

An attractive feature of Hansen's (1999) approach in comparison to Andrews and Chen (1994) is that the grid bootstrap method does not impose the assumption that the shocks are i.i.d. Gaussian. This feature is particularly appealing to the current study because in practice the null hypothesis of normally distributed shocks is generally rejected at the 1 percent level of significance. In addition, Hansen's (1999) grid bootstrap approach provides a confidence interval with correct asymptotic coverage for both stationary and local-to-unity AR models. Therefore, in the following full sample and rolling estimations, we focus on the grid bootstrap estimates for inflation persistence, while the OLS and Andrews-Chen MU estimates are provided for comparison.

Another important issue for the AR model specification is the selection of the optimal lag order. We set the optimal lag order based on Akaike

Information Criterion (AIC) with a maximum of eight lags for both inflation series over the underlying sample periods. In practice, this setup appears to balance well between (desirable) parsimony and (undesirable) serial correlation.

In order to investigate the structural stability of inflation persistence, we will start by estimating model (1) for the China's inflation series over the entire sample period and rolling samples of a 10-year window. By design, the full sample estimation produces benchmark values under the assumption of model stability, while the rolling-sample estimation allows the underlying model to be flexible over the specified moving windows. By inspecting a sequence of the rolling estimates and comparing the estimates across different windows and with the full sample results, we provide preliminary evidence of variation in inflation persistence for China.

### 2.3. Full Sample Estimation

Based on the preceding setup, Table 1 reports full sample estimates of inflation persistence using the grid bootstrap, the Andrews-Chen, and the OLS estimator for the two inflation series. In these regressions, heteroskedasticity consistent covariance matrix estimator (HCCME) with finite-sample degree of freedom correction is used; the  $p$ -values of the serial correlation test (denoted  $p$ -auto) are based on  $F$  rather than  $\chi^2$  statistic for the Breusch-Godfrey type of LM test due to Kiviet (1986).

**TABLE 1.**

Full sample estimates of inflation persistence: 1981Q1-2007Q1

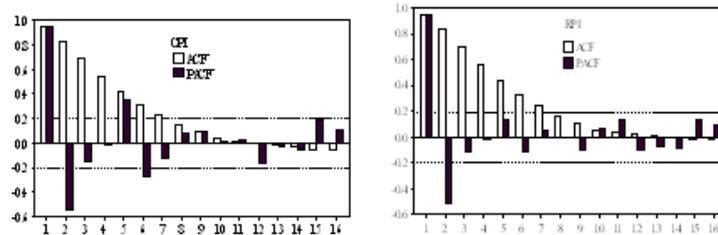
	grid bootstrap		Andrews-Chen		OLS				
	MU	90% interval	MU	90% interval	$\hat{c}$	$\hat{\rho}$	$\bar{R}^2$	$p$ -auto	lag
CPI	0.940	[0.882, 1.011]	0.939	[0.866, 0.963]	0.49 (0.20)	0.922 (0.03)	0.94	0.56	6
RPI	0.942	[0.866, 1.019]	0.933	[0.863, 0.958]	0.42 (0.19)	0.918 (0.04)	0.92	0.35	2

Notes: MU refers to median unbiased estimate;  $p$ -auto refers to  $p$ -value of Breusch-Godfrey LM test for serial correlation up to 4 lags; lag denotes optimal lag order chosen by AIC. The grid bootstrap estimates were computed using Hansen's (1999) method with 200 gridpoints and 1999 bootstrap replications, while Andrews-Chen estimates were computed using Andrews and Chen (1994) method with 1000 Monte Carlo simulations. The innovations for the simulated processes were based on drawing from the estimated residuals. In the OLS regressions, heteroskedasticity-robust standard errors are reported in parentheses.

The following points should also be noted in relation to Table 1. First, in all the regressions, the optimal lag lengths specified by AIC are sufficient to rule out serial correlation in model (1). Second, the sample partial autocorrelation functions (PACF) together with the sample autocorrelation

function (ACF) for the inflation series over the full sample are plotted in Figure 2. They show that the corresponding PACF for each inflation series cuts off abruptly to zero (within the 2 standard error bounds) after the corresponding optimal lag orders chosen by AIC, indicating that the  $AR(p)$  models specified here are indeed plausible. Third, the OLS estimates of the persistence parameter for the CPI and RPI inflation are quite similar in magnitude at around 0.92.

**FIG. 2.** Sample ACF and PACF of China CPI and RPI inflation



Notes: Two horizontal dotted lines in each graph denote the approximate two standard error bounds associated with the PACF.

In addition and as expected, the MU estimates of both the grid bootstrap and the Andrews-Chen methods are higher than the corresponding (biased) OLS point estimates. In both cases, however, the bias correction appears to be small (around 2 percent). The grid bootstrap 90 percent confidence interval generally includes the value of unity, which suggests that the AR process for the inflation series may contain a unit root.

Taken as whole, the results in Table 1 suggest that both the CPI and RPI inflation are highly persistent under the assumption of no structural shifts in the dynamic process of the AR model.

#### 2.4. Rolling Estimation

To assess whether the preceding suggestion holds across time and to examine the extent to which there have been changes in the process generating the underlying inflation series, we present in Figure 3 plots of rolling estimates (10-year rolling window) of  $\rho$  using the OLS, the grid bootstrap, and the Andrews-Chen procedures respectively.

By construction, each estimate relates to a model of the form (1) with specific lag length chosen by AIC for each sequential sample. If the persistence parameter remains stable, we would expect a sequence of largely time-invariant estimates. Hence, the rolling estimates can provide preliminary evidence on the potential structural shifts in the coefficient of our interest.

From Figure 3, we observe that both the OLS and the grid bootstrap median unbiased rolling estimates show some variations in inflation persistence over time. In particular, there is a dramatic downward shift in the degree of persistence of the CPI inflation in the late 1990s. Indeed, the estimates of  $\rho$  for the CPI inflation exhibit a transitory drop in 1993, returning to its high level in 1994, then appears to decline dramatically around 1995, and has become much lower since then.

Interestingly, however, the estimates of the persistence parameter for RPI inflation also appears to become lower in the middle 1990s, but returns back to a higher level quickly. Further investigation of Figure 3 also reveals that the rolling estimates of the Andrews-Chen method only witness a sudden increase in 1994 (and a second rise at the end of the rolling sample for the RPI) but the increase seems transitory in nature.

Taken as a whole, the preceding results suggest that the high inflation persistence obtained over the full sample may not be invariant over time, in particular for the CPI series. This prompts us into formal structural break tests for inflation persistence in the AR model.

### 3. UNKNOWN STRUCTURAL BREAK TESTS

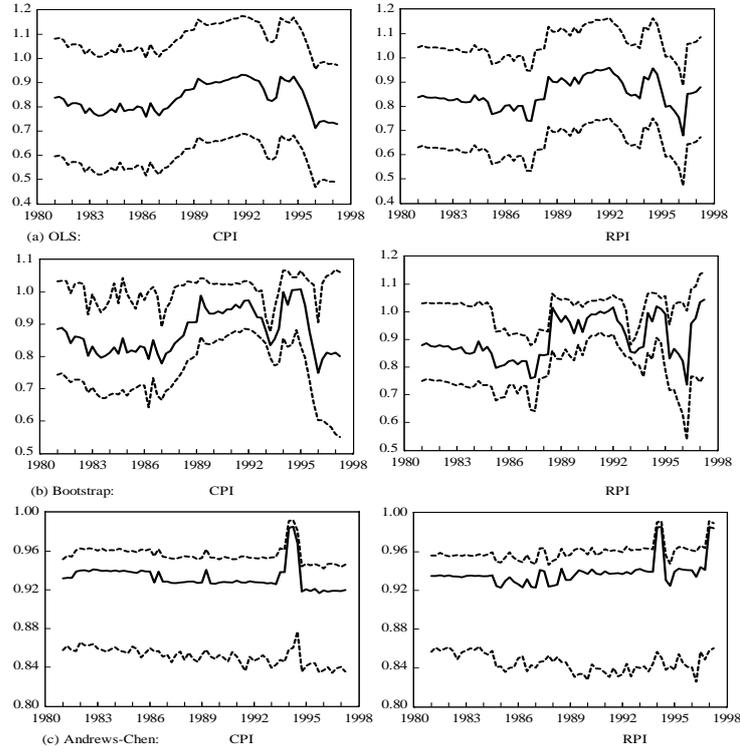
The rolling estimates in the previous section provide an intuitive impression of potential structural shifts over time in the persistence parameter. However, that approach does not allow us to draw firm conclusions about the significance of the structural breaks. Therefore, in this section we perform formal tests to investigate the statistical nature and the timing of potential structural breaks in model (1) that focus on the persistence parameter.

Theoretical advances in the literature of unknown structural break tests, in particular the important contributions by Andrews (1993), Andrews and Ploberger (1994) and Hansen (1997), enable us to identify changes and the associated timing in the underlying model with considerable precision. In this paper, we employ the Supreme Wald test of Andrews (1993) and the Exponential- and Average- Wald tests of Andrews and Ploberger (1994) to test for unknown structural breaks in the univariate AR process of Chinese inflation. All three tests are designed to test for the same null hypothesis of no structural break in the underlying parameters of interest. The corresponding  $p$ -values of these underlying tests are computed using the method of Hansen (1997).

By construction, the Andrews' (1993) Supreme Wald statistic is the maximum Wald-statistic for testing a break through all possible break points over a specified searching range, say  $\tau$ , which is given by

$$SupWald = \sup W_T(\tau_i) \mid \tau_i \in [\tau_{\min}, \tau_{\max}] \quad (3)$$

**FIG. 3.** Rolling estimates of inflation persistence



Notes: In each graph, the solid line represents median unbiased estimates (panel (a) for OLS point estimates) for the persistence parameter in the AR process over 10-year rolling windows starting from the sample of 1981Q1-1990Q4. The dotted lines denote two standard error bounds associated with the persistence parameter estimates.

where  $W_T(\tau_i)$  denotes the sequential Wald-statistic testing for the null hypothesis of no structural break in the underlying parameter. We set a customary searching interval  $\tau_i \in [0.15, 0.85]$  of the full sample  $T$  to allow a minimum of 15 percent of effective observations contained in both pre- and post-break periods.

In addition, the Andrews and Ploberger's (1994) average- and exponential-statistics are given by

$$AveWald = \int_{\tau_{\min}}^{\tau_{\max}} W_T(\tau) d\tau, \quad (4)$$

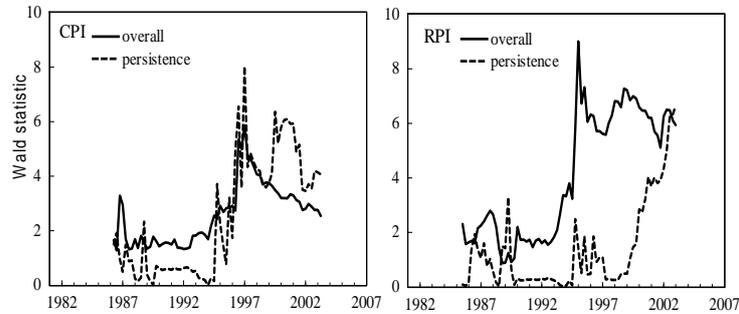
and

$$ExpWald = \ln \left\{ \int_{\tau_{\min}}^{\tau_{\max}} \exp[0.5W_T(\tau)] d\tau \right\}. \quad (5)$$

In practice, HCCME versions of the statistics are computed throughout the tests based on the residuals under the null hypothesis of no structural break for computational convenience.

A plot of the estimated sequential Wald-statistic provides information about the timing of a structural break occurring in the underlying parameters. Figure 4 presents the estimates of the sequential Wald-statistic testing for constancy of the overall parameters and the persistence parameter in the AR process for the CPI and RPI inflation series respectively. The graphs show that a maximum Wald statistic is obtained around 1997 for both the overall parameters and persistence parameter associated with the CPI inflation, while for the RPI inflation a supreme Wald statistic occurs in 1995 for the overall parameters and 2003 for the persistence parameter<sup>1</sup>.

**FIG. 4.** Sequential Wald-statistics of structural break tests



Notes: Overall and persistence refer to sequential Wald-statistics of testing structural change in all parameters and persistence parameter of the AR model, respectively.

The numerical estimation results are summarized in Table 2, where for each inflation series, stability tests are performed on all the coefficients overall and the three individual coefficients. The Table reports  $p$ -values of the Andrews-Ploberger family of statistics (denoted  $p$ -sup,  $p$ -exp, and  $p$ -ave) for the structural break tests over full sample of 1981Q1-2007Q1.

The first statistic (overall) tests for stability of all the coefficients in model (1). The second statistic ( $c$ ) tests for stability of the intercept term

<sup>1</sup>Note that the break in 2003 corresponding to the result for persistence parameter of the RPI may be unreliable since the break occurs at an extreme of the searching interval.

**TABLE 2.**

Andrews-Ploberger structural break tests for AR model

Inflation	Coef.	$p$ -sup	$p$ -exp	$p$ -ave	Break-date
CPI	overall	0.000	0.000	0.001	1996Q4
	$c$	0.035	0.024	0.046	1994Q3
	$\rho$	0.065	0.062	0.069	1996Q4
	$\alpha_{\Delta}$	0.042	0.027	0.044	1999Q2
RPI	overall	0.000	0.001	0.000	1994Q4
	$c$	0.012	0.019	0.071	1994Q3
	$\rho$	0.118	0.121	0.191	2003Q1
	$\alpha_{\Delta}$	0.000	0.000	0.000	1998Q3

Notes: Sample spans 1981Q1-2007Q1 prior to lag adjustment.  $p$ -sup,  $p$ -exe and  $p$ -ave denote  $p$ -values of the Supreme-, Exponential- and Average-Wald tests in Andrews (1993) and Andrews and Ploberger (1994), which were computed using the method of Hansen (1997); Break-date refers to structural break date in which the Supreme-Wald statistic is obtained. The searching interval of the tests is central 70% of the sample.

assuming that the remaining coefficients are constant. Likewise, the third statistic ( $\rho$ ) tests for constancy of the persistence parameter assuming the other coefficients are constant and finally, the last statistic ( $\alpha_{\Delta}$ ), tests for joint stability of the dynamic coefficients (i.e. the coefficients on all first difference inflation terms in the model) assuming the intercept and persistence parameter are constant. For each regression, the optimal lag order in the AR model is chosen by the AIC with maximum eight lags.

The structural break tests for the overall parameters in Table 2 indicate statistically significant evidence of instability in both inflation series, with  $p$ -values of all the Andrews-Ploberger family of statistics smaller than 1 percent. Investigating the other three tests in each regression reveals that the instability for the CPI inflation is primarily concentrated in the persistence parameter while the instability for the RPI inflation appears to originate from the intercept and the dynamic coefficients  $\alpha_{\Delta}$ .

Overall, the Andrews-Ploberger statistics in Table 2 suggest that in general, the AR process of both the CPI and RPI inflation is unstable over 1981-2007, but the origins of the instability in the dynamic process for the two inflation series appears to be different from each other. In particular, we find a significant structural change in the persistence parameter of the CPI inflation around 1997. The persistence parameter in the AR model for the RPI inflation does not manifest statistically significant change. This disparity may reflect the conceptual differences between the CPI and RPI inflation series. For instance, the market focus and commodity baskets between the two inflation series differ from each other; more importantly,

the impact of price regulations on the RPI in China is limited since the components of the RPI do not include service items<sup>2</sup>.

#### 4. INFLATION PERSISTENCE OVER THE IDENTIFIED REGIMES

Based on the results and discussions in the preceding section, we now proceed to investigate the nature of China inflation persistence over the identified regimes. Specifically, we estimate the persistence parameter of inflation in the dynamic process of model (1) over different regimes separated by the structural breaks presented in Table 2. For instance, the persistence of the CPI inflation is estimated over two sub sample periods segmented by the structural break point of 1997Q1.

Although the structural break tests in the previous section indicate that the persistence parameter of the RPI inflation has not changed significantly, the overall parameters in the AR model for the RPI inflation does manifest a significant break. Thus we also provide the estimates of the RPI inflation persistence before and after its overall break point.

For comparison purposes, the MU estimates of Hansen's grid bootstrap method, the Andrews-Chen method and the OLS estimate of inflation persistence over the identified regimes are all reported. The estimation results for both inflation series are summarized in Table 3.

From Table 3, it is evident that the statistical nature of the persistence parameter for the CPI inflation before and after 1997 indeed differs markedly. For example, according to the grid bootstrap median unbiased estimates, the persistence of the CPI inflation is high (MU of  $\hat{\rho} = 0.890$ ) over 1981-1996, and comparatively low (MU of  $\hat{\rho} = 0.814$ ) over 1997-2007. The MU estimates of Andrews-Chen and the OLS estimates of the persistence parameter for the CPI inflation provides a more striking difference between the two regimes. For instance, the Andrews-Chen estimate of CPI inflation persistence is 0.898 before 1997 while the magnitude drops to 0.795 after the structural break date. The OLS estimates of CPI inflation persistence before and after 1997 are also striking (0.868 vs 0.745).

For the RPI inflation, the estimates of the persistence parameter after its break date (1995) using the alternative estimation methods is smaller than that before the break date, albeit the magnitude of the estimates of the persistence parameter in both periods is comparatively higher than that of the CPI.

Another notable result is that, for both of the inflation series, the  $\bar{R}^2$  is around 0.90 in the regressions before the respective break points while

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<sup>2</sup>We leave the investigation of these possible explanations pertaining to the different results here to future research.

it drops considerably after the breaks. This may reflect the fact that the variation of the inflation series before the break point is much more evident than after the break point, as attested in Figure 1. Thus the estimated standard deviation over the period of less variation (after break point) is inevitably small. This intuition may also explain why the 90% confidence intervals of both the grid bootstrap and the Andrews-Chen methods after the break are wider than the corresponding ones before the break.

To assess the sensitivity of our findings, we examine whether the underlying results are robust to slight changes of the break point. For example, we estimate inflation persistence for the CPI inflation over several sub samples prior to 1997, namely 1981Q1-1996Q3, 1981Q1-1996Q2, and 1981Q1-1996Q1, and examine whether the estimates of inflation persistence over these sub samples are indeed high. Likewise, we also estimate the persistence parameter over several sub samples post 1997 and compare the results with those for the samples before 1997. The corresponding results (in Table 3) suggest that our baseline findings do not alter in this experiment.

## 5. SENSITIVITY ANALYSIS

In our preceding investigations, we have been focusing on estimating the persistence parameter in the univariate inflation process, in which we may have effectively squeezed a potential real driving variable into the error term. However, the exclusion of any real driving variables may result in spurious findings of a high value for the inflation persistence parameter owing to the omission of possible feedback from the real variable to inflation.

In addition, O'Reilly and Whelan (2005) show that unless one conditions the dynamic regression of inflation on appropriate driving variables, it is inappropriate to make any direct link between the persistence parameter of inflation and the true effect on current inflation of its own lagged observations.

Therefore, we investigate the robustness of our baseline finding by taking into account the impact of real economic growth on inflation. To achieve this, we augment model (1) by adding a measure of real economic growth and estimate the persistence parameter in a conventional Phillips curve specification (e.g. Stock and Watson, 1999), viz.

$$\pi_t = c + \rho\pi_{t-1} + \sum_{i=1}^{p-1} \alpha_{\Delta i} \Delta \pi_{t-i} + \beta y_{t-1} + u_t \quad (6)$$

**TABLE 3.**

Sub sample estimates of China inflation persistence in AR model

CPI	Sub Samples	Grid bootstrap		Andrews-Chen		OLS				
		MU	90% interval	MU	90% interval	$c$	$\hat{\rho}$	$\bar{R}^2$	$p$ -auto	lag
Pre-break	1981Q1-1996Q4	0.890	[0.813, 0.982]	0.898	[0.779, 0.937]	1.290 (0.453)	0.868 (0.043)	0.92	0.91	6
	1981Q1-1996Q3	0.889	[0.812, 0.981]	0.894	[0.781, 0.932]	1.288 (0.454)	0.867 (0.042)	0.92	0.92	6
	1981Q1-1996Q2	0.894	[0.815, 0.987]	0.902	[0.774, 0.940]	1.276 (0.454)	0.871 (0.042)	0.92	0.95	6
	1981Q1-1996Q1	0.897	[0.815, 1.002]	0.899	[0.749, 0.937]	1.276 (0.456)	0.872 (0.044)	0.92	0.95	6
Post-break	1997Q1-2007Q1	0.814	[0.569, 1.067]	0.795	[0.513, 0.863]	0.307 (0.144)	0.745 (0.124)	0.74	0.90	4
	1997Q2-2007Q1	0.820	[0.569, 1.066]	0.773	[0.520, 0.851]	0.294 (0.143)	0.729 (0.124)	0.75	0.76	4
	1997Q3-2007Q1	0.804	[0.548, 1.060]	0.801	[0.539, 0.873]	1.276 (0.151)	0.745 (0.121)	0.75	0.69	4
	1997Q4-2007Q1	0.811	[0.567, 1.070]	0.803	[0.489, 0.878]	0.271 (0.161)	0.743 (0.122)	0.74	0.74	4
RPI										
Pre-break	1981Q1-1994Q4	1.008	[0.799, 1.047]	0.925	[0.796, 0.952]	1.044 (0.457)	0.895 (0.072)	0.88	0.94	2
	1981Q1-1994Q3	1.008	[0.800, 1.046]	0.888	[0.758, 0.927]	1.176 (0.433)	0.861 (0.064)	0.88	0.88	2
	1981Q1-1994Q2	0.899	[0.776, 1.032]	0.891	[0.761, 0.925]	1.46 (0.440)	0.869 (0.067)	0.88	0.90	2
	1981Q1-1994Q1	0.910	[0.777, 1.035]	0.869	[0.738, 0.915]	1.192 (0.427)	0.845 (0.061)	0.89	0.93	2
Post-break	1995Q1-2007Q1	0.895	[0.767, 1.025]	0.907	[0.609, 0.944]	-0.017 (0.180)	0.845 (0.064)	0.82	0.54	6
	1995Q2-2007Q1	0.896	[0.771, 1.025]	0.880	[0.583, 0.925]	-0.056 (0.174)	0.822 (0.067)	0.82	0.62	5
	1995Q3-2007Q1	0.886	[0.753, 1.027]	0.872	[0.604, 0.918]	-0.056 (0.174)	0.810 (0.068)	0.79	0.59	5
	1995Q4-2007Q1	0.853	[0.725, 1.008]	0.856	[0.563, 0.917]	-0.062 (0.175)	0.797 (0.064)	0.76	0.23	5

where  $y$  denotes the rate of real GDP growth<sup>3</sup>.

<sup>3</sup>The quarterly data for the (year-on-year) growth rate of the real GDP is computed based on the level of real GDP published by the National Bureau of Statistics of China.

Based on the augmented model (6), we first perform the Andrews-Ploberger structural break tests for the model and the results are shown in Table 4. These results indicate that although the timing of the breaks for the individual parameters are not exactly the same as those shown in Table 2, the break dates for the overall parameter stability of model (6) coincides with those associated with model (1). In addition, the persistence parameter for the CPI inflation in the augmented specification manifests a significant structural change (the corresponding  $p$ -values of the structural break tests are smaller than 5%) while the RPI inflation persistence parameter experiences no significant structural break at 5% level of significance.

Since our primary interest is the changes in inflation persistence, Table 5 then provides sub sample estimation results of the augmented model for both the CPI and RPI inflation series, using both the grid bootstrap and the OLS estimators<sup>4</sup>. The first result worth noting is that, while admittedly crude, the measure of the real economic growth plays a statistically significant role in the augmented specification in virtually all regressions (except for the RPI post-break regression in Table 5).

Moreover, the reduction of the CPI inflation persistence after 1997 in model (6) is more pronounced than that shown in Table 2, using either the grid bootstrap method or the traditional OLS estimator. For example, the MU estimate for the CPI inflation persistence is 0.917 before 1997 while the estimate declines to 0.632 after the break date. The OLS estimate of the persistence parameter for CPI inflation is 0.889 prior to 1997 and the estimate drops considerably to 0.551. There is also a sizable change in RPI inflation persistence, although the change is less striking: the difference between the two sample periods is around 0.12.

Taken as whole, the general pattern of the structural changes in inflation persistence in the augmented model does resemble that in the AR process studied in the previous section. In practice, we also investigated the sensitivity of the baseline findings when the timing of the structural break in the intercept of the model is allowed to vary and a possible shift in the dynamic coefficients is considered. In general, however, there are no substantive changes for the findings in Table 5, suggesting the robustness of the baseline findings.

## 6. POLICY IMPLICATIONS

The empirical results in the foregoing sections suggest that the persistence of inflation in China has witnessed declines since the end of the 1990s,

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<sup>4</sup>The rate of real GDP growth is treated as fixed regressor in the grid bootstrap estimation. Since the Andrews-Chen method is specifically designed for the dynamic univariate process, the corresponding results are not reported.

**TABLE 4.**

Andrews-Ploberger structural break tests for the augmented model

Inflation	Coef.	$p$ -sup	$p$ -exp	$p$ -ave	Break-date
CPI	overall	0.000	0.000	0.000	1996Q4
	$c$	0.045	0.063	0.097	1994Q3
	$\rho$	0.014	0.012	0.024	1986Q1
	$\alpha_\Delta$	0.002	0.002	0.001	1996Q2
	$\beta$	0.021	0.034	0.067	1994Q3
RPI	overall	0.000	0.000	0.001	1994Q4
	$c$	0.008	0.014	0.065	1994Q3
	$\rho$	0.088	0.052	0.066	2003Q1
	$\alpha_\Delta$	0.004	0.001	0.003	1998Q1
	$\beta$	0.008	0.014	0.069	1994Q3

**TABLE 5.**

Estimation results of the augmented model

	grid bootstrap		OLS					
	MU	90% interval	$\hat{c}$	$\hat{\rho}$	$\hat{\beta}$	$\bar{R}^2$	$p$ -auto	lag
CPI								
Full sample	0.935	[0.882, 1.003]	-1.358 (0.675)	0.917 (0.029)	0.186 (0.029)	0.94	0.89	6
Pre-break	0.917	[0.843, 1.006]	-0.674 (0.930)	0.889 (0.039)	0.166 (0.080)	0.93	0.98	6
Post-break	0.632	[0.338, 1.049]	-3.077 (1.215)	0.551 (0.151)	0.391 (0.143)	0.79	0.26	4
RPI								
Full sample	0.925	[0.855, 1.012]	-1.341 (1.657)	0.900 (0.039)	0.187 (0.072)	0.93	0.49	2
Pre-break	0.943	[0.816, 1.037]	-0.825 (0.878)	0.885 (0.063)	0.190 (0.082)	0.90	0.98	2
Post-break	0.819	[0.641, 1.025]	-1.298 (1.670)	0.793 (0.104)	0.140 (0.179)	0.82	0.06	6

although the reduction in the RPI inflation is less striking. The reduction in inflation persistence has two important implications. One is related to the low inflation environment that China has experienced since the end of the 1990s and the other is associated with the remarkable phenomenon of high economic growth under the low inflation environment in China during the past decade.

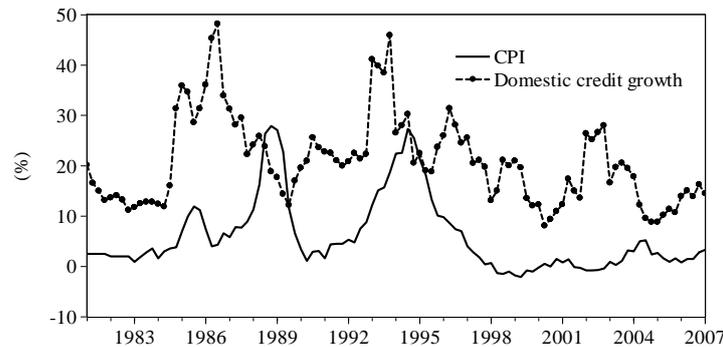
First, the reduction in inflation persistence has helped entrench low inflation. As discussed in the outset, if the inflation process has become less

persistent, shocks that trigger inflation now would have a less protracted influence on the rate of inflation in the ensuing period. Hence the structural change of inflation persistence can be viewed as an important factor that has helped maintain low inflation environment in China over the period since the end of the 1990s.

This assertion highlights the importance of macro and monetary policy changes in reducing inflation since the early 1990s, since it implies the important dynamic interaction between the policy changes, the persistence of inflation, and the rate of inflation contribute to increased stability.

More specifically, it is widely agreed that the growing concern of the Chinese government about prevailing high inflation of the early 1990s has motivated the central government to implement a number of macro and monetary policy adjustments. One notable change is the more restrictive credit policy. To see this more clearly, Figure 5 plots the time series of the rate of domestic credit growth in conjunction with the CPI inflation in China over 1981Q1-2007Q1 and shows how the high peaks of credit expansion have preceded the high peaks of CPI inflation.

**FIG. 5.** CPI inflation and domestic credit growth in China



Data Source: International Financial Statistics.

It can be seen from Figure 5 that the rate of domestic credit growth in 1993 is above 45 percent, while it drops to about 30 percent in 1994. Except for a few local peaks after 1995, the domestic credit growth after the middle 1990s exhibit a generally decreasing pattern, which corresponds to the restrictive policy that China has implemented over the period.

The important policy change may have induced structural instability of the dynamic process of inflation, in particular the persistence of inflation. Of course, other policies including stabilizing financial markets, strengthening the regulation and supervision on financial institutions, and restricting

investment in fixed assets, may have also contributed to the change in inflation persistence.

In essence, the fact that the change in macro policies induces the change in inflation persistence is consistent with the well-known Lucas Critique (Lucas, 1976). The Lucas Critique suggests that reduced forms of macroeconomic models, particularly lagged autoregressive models are unlikely to be stable given policy shifts over time. Thus it may be plausible to infer that the notable policy changes in China in the 1990s cause a significant structural change in inflation persistence which in turn helps to keep inflation at a low level.

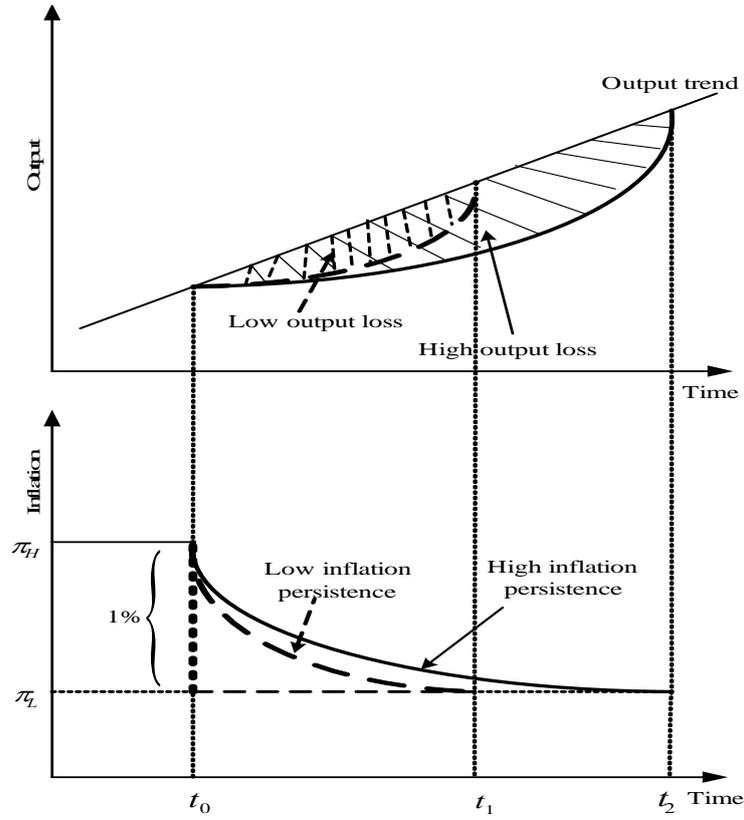
Second, the finding of the change in China inflation persistence has important implications for understanding the seemingly puzzling fact about China's experience over the past 10 years, viz. that high growth has been accompanied with low inflation. As discussed in Fuhrer (1995), a high degree of inflation persistence implies that disinflation following an adverse shock entails higher transitory output costs. In addition, inflation expectations may be difficult to be anchored during the period of high inflation persistence (Melick and Galati, 2006). Consequently, implementation and communication of stabilizing macro and monetary policy become more difficult in the regime of highly persistent inflation.

In contrast, during the period of low inflation persistence, the cost of disinflation becomes much lower. This is illustrated by Figure 6, which shows possible time paths of real output and inflation in response to a tighter monetary policy that ultimately lowers inflation by one percentage point, under different inflation persistence. Apparently, high inflation persistence induces high output loss (i.e. low economic growth) while low inflation persistence corresponds to low output loss.

Hence, in the case of low inflation persistence, a country can keep its output growth at a relatively high standard with the rate of inflation being at a low level. Figure 7 depicts the rate of the real GDP growth and the CPI inflation over 1981Q1-2007Q1 in China, which reinforces this reasoning. Figure 7 shows that the growth rate of the real GDP in China witnessed a dramatic decline along with the large reduction of the CPI inflation during the anti-inflation wave over the first half of the 1990s. In particular, the real GDP growth rate dropped to about 7 percent by the end of 1997. Nevertheless, real economic growth started increasing steadily after 1997 and it has fluctuated around 10 percent since the early 2000s. This reflects the fact that real disinflation cost in China has been relatively low during the period of low inflation persistence.

Therefore, the reduced CPI inflation persistence in China effectively helps maintain inflation at a relatively low level during the period of high economic growth. This may in large part, if not all, explain why China has

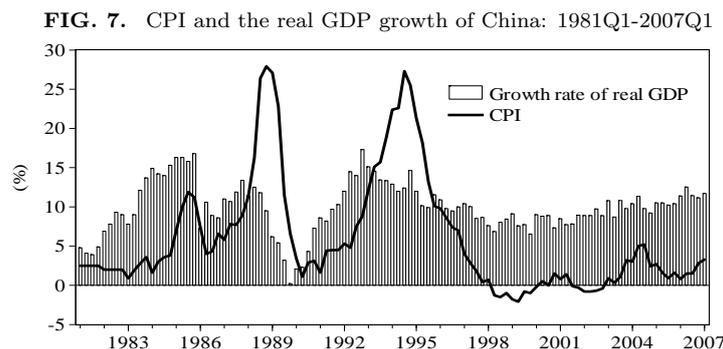
FIG. 6. Output loss associated with different inflation persistence



had a high rate of economic growth but low inflation since the end of the last century.

### 7. CONCLUSIONS

This paper investigates structural instability in China inflation persistence over the most recent quarter of a century. Employing the recently developed unknown structural break tests in conjunction with two median unbiased estimation approaches, we show that there is a significant reduction in the persistence of consumer price inflation. Although the persistence of retail price inflation has not experienced a statistically significant break, it also manifests some reduction. We associate this finding with profound policy changes in China during the period of early 1990s. In particular, we propose that policy changes in China credit expansion have induced a



Data source: National Bureau of Statistics of China and International Financial Statistics.

structural break in the CPI inflation persistence, through which the inflation has become entrenched at a low level.

This finding is in broad agreement with Guerineau and Guillaumont (2005) who also underscores the importance of policy changes in dampening inflation in China. Nonetheless, our finding offers a deeper insight and provides a new perspective for us to pinpoint the specific pattern of high economic growth accompanied with low inflation in China over the most recent decade.

In addition, the empirical results suggest that the statistical nature of the structural changes in the persistence of the consumer price inflation and the retail price inflation is different. In particular, the source of structural shifts in the dynamic models for the RPI inflation appears to concentrate in the intercept and dynamic coefficients rather than the persistence parameter. This finding may reflect the important impact of limitations of price regulations on services in China. The finding may also imply the differences of the influences of macro and monetary policies on the CPI inflation and the RPI inflation. This last proposal is of course informal. More rigorous study of the relevant issue would entail a multivariate dynamic model in conjunction with counterfactual simulation analysis. Stock and Watson (2002) and Estrella and Fuhrer (2003) provide important contributions in this direction for the US, and future research may undertake this task for the relevant issue of China.

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