Inflation persistence, inflation expectations, and monetary policy in China

Chengsi Zhang *

China Financial Policy Research Center, School of Finance, Renmin University of China, No. 59 Zhong Guan Cun Street, Haidian District, Beijing, 100872 China

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This paper constructs a quarterly series of GDP deflator inflation for China from 1979 to 2009 and tests for a structural break with an unknown change point in the dynamic inflation process. Empirical results suggest a significant structural change in inflation persistence. Employing a counterfactual simulation method, we show that the structural change is primarily attributed to better conduct of monetary policy and the resultant better anchored inflation expectations. This finding implies that the quiescence of inflation in China over the past decade could well be followed by a return to a high inflation era in the absence of a determined effort by the monetary authorities in managing inflation expectations. Therefore, the use of a preemptive monetary policy to anchor inflationary expectations and to keep inflation moderate is warranted in China.

C22

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1. Introduction

Since 1979, China has witnessed remarkable changes in the pattern of its inflation process. In particular, the rate of inflation in China is now much lower and less volatile than it was in the 1980s and early 1990s. This change coincides with China’s monetary policy reforms since the late 1990s. Significant examples of these reforms include establishments of interbank money markets and bond markets in 1995–1996 and commencement of open market operations in 1998, through which China has changed its monetary policy from direct credit quota control to indirect adjustment with both quantity- and price-based instruments.

These shifts in monetary policy may have induced major structural changes in the dynamic models of the inflation process, as articulated in the famous Lucas critique (Lucas, 1976). In the context of the Chinese economy, these important changes in monetary policy may have induced structural instability in the dynamic process of inflation, and in particular the persistence of inflation. That is, the change in inflation dynamics could reflect the fundamental shift in monetary policy whereby the PBOC now systematically acts to stabilize inflation around a potential long-run target and has gained credibility with the public that it will continue to do so into the future.

Surprisingly, relatively little effort seems to have been devoted to testing the empirical importance of the Lucas critique in terms of China’s inflation dynamics and monetary policy. The existing literature focuses particularly on issues relevant in the United States, while little consensus has been achieved. For instance, Estrella and Fuhrer (2003) and Rudebusch (2005) assess the empirical importance of the Lucas critique in US monetary models. Their results suggest that the inflation process with backward-looking descriptions of expectation formation experiences no structural change even when monetary policy changed. However, recent studies by Kim and Kim (2008), Zhang et al. (2008), and Zhang and Clovis (2009a) suggest a link between changes in inflation dynamics and monetary policy shifts, indicating the empirical importance of the Lucas critique.

Notwithstanding these empirical controversies, standard monetary models with inflation dynamics for China assume that the monetary policy regime in China is fixed. For example, Hasan (1999) argues that there is a stable feedback relationship between inflation and monetary growth in China. Guerineau and Guillamont (2005) estimate a dynamic equation of consumer price variations for China over the period 1986–2002, assuming no structural change in monetary policy and the underlying model for inflation dynamics. Zhang (2009) attempts to specify and estimate a price-based (i.e. interest rate) monetary policy reaction function, whilst assuming that monetary policy in China has experienced no structural shift. Recently, Zhang and Clovis (2009b) noted possible changes in China’s monetary policy regime, and the parallel change in inflation dynamics. The focus there, however, is on consumer price inflation,
and no quantitative evidence linking the change in inflation process and the shift in monetary policy is provided.

This paper considers a different and yet (hopefully) more important measure of price inflation in China, namely Gross Domestic Product (GDP) deflator inflation (denoted GDPDI), spanning a relatively long period of time, from 1979 to 2009. We construct a quarterly series of GDPDI by using published data from China’s National Bureau of Statistics (NBS) for the growth rate of real GDP and the level of nominal GDP, in conjunction with the estimation method in Abeysinghe and Rajaguru (2004). We then test for structural change in the dynamic models of the inflation process, using structural break tests at unknown break point.

In addition, to quantify the link between the structural break in inflation dynamics and monetary policy changes, we specify a multivariate dynamic model to capture dynamic interactions between the real economy, inflation, and monetary policy, and then work through counterfactual simulations. Our empirical results suggest that the structural change in the dynamic models of the inflation process is attributed to improvements in the conduct of China’s monetary policy and the resultant better anchored inflation expectations.

To this end, Section 2 describes the construction of the GDP deflator inflation series. Section 3 specifies alternative models for inflation dynamics and identifies possible structural break in the underlying models. Section 4 documents systematic changes in monetary policy and discusses the corresponding changes in inflation expectations. Section 5 then performs counterfactual simulations to quantify the contribution of monetary policy shifts to the change in inflation persistence. Section 6 concludes the paper.

2. Data descriptions

The inflation series considered in this paper are chosen to provide relations that are of most relevance for monetary policy analysis and to represent general inflation in an overall economy. The rate of inflation based on the consumer price index (CPI) appears to be the most closely watched economic barometer in China. Yet in recent years, as the public and the media have paid considerable attention to monthly CPI announcements by China’s NBS, more and more economists as well as the public have expressed their dissatisfaction with the way the CPI measures inflation in China, especially in its treatment of the cost of housing.

The GDP deflator, by contrast, is a broader measure of price changes than the CPI, as it averages prices of all goods and services in the economy. Since the “market basket” of the GDP deflator is essentially the whole production of the nation, the GDP deflator inflation reflects general price changes in the overall economy more accurately. Therefore, this paper aims to characterize the nature of China’s inflation dynamics in terms of the broader measure, namely the GDP deflator inflation.

Since there are no published data available for the GDP deflator series with quarterly frequency, we need to first construct quarterly data for the GDP deflator. Note that we do not use annual data because with annual data most economic relationships are likely to become simply contemporaneous due to temporal aggregation. In China, quarterly data of nominal GDP in levels and real GDP in growth rates (year-on-year) are available from the NBS since 1992Q1. These are used to construct a quarterly real GDP series with 1997 as the base year. Using the quarterly nominal and real GDP series we can derive the GDP deflator for 1992Q1–2009Q4.

Prior to 1992, however, China’s GDP data (both levels and growth rates) are mostly available only on an annual basis. To address the problem of low frequency, we first convert the annual data of the nominal GDP over 1978–1991 (published by the NBS) into quarterly data by averaging annual figures, and then employing estimation results from Abeysinghe and Rajaguru (2004) for growth rates of real GDP (year-on-year) over 1978Q1–1991Q4 to derive the quarterly real GDP series. The nominal and real GDP series of quarterly frequency are then used to obtain the GDP deflator series.

Note that the methodology in Abeysinghe and Rajaguru (2004) is effectively the Chow–Lin estimation method in the spirit of Chow and Lin (1971). The basic idea is to find the GDP-related quarterly series and derive a forecasting equation by running a regression of annual GDP on annual related series. Quarterly figures of the related series are used to forecast the quarterly GDP series which are adjusted to match the annual aggregates. The estimated quarterly real GDP series based on this method appears to match the officially published annual data quite convincingly.

Using the GDP deflator series over 1978Q1–2009Q4, we then calculate the year-on-year growth rate of the underlying series and obtain the corresponding GDP deflator inflation (i.e. \( \pi_t = (P_t/P_{t-1} - 1) \times 100 \)), where \( \pi_t \) denotes GDP deflator inflation and \( P_t \) denotes the GDP deflator series. The resultant quarterly GDP inflation series is depicted in Fig. 1.

The plot of the time series data in Fig. 1 suggests that since the commencement of economic reforms in the late 1970s, inflation in China witnessed its first distinct increase during the period 1979–1981, followed by spikes in 1983 and 1985. The two most striking peaks occurred in the late 1980s and the middle 1990s. Since the late 1990s, however, inflation in China has been relatively low and stable, even witnessing three years of deflation (1997–2000). Despite the two notable peaks of inflation occurring in 2004–2005 and 2007–2008 due to transitory demand shocks (e.g. shock to real estate market) and supply shocks (e.g. shock to food and energy prices) respectively, the most recent decade may be characterized as a low inflation era.

The evolution of the GDP deflator inflation over the past three decades reflects the corresponding historical changes in the mechanism of price formation in China, accompanied by changes in monetary policies. To be specific, it is well known that prices of most commodities in China were administered by government agencies and changed infrequently until the end of the 1970s. Since the start of economic reforms in 1978, however, the government-set prices were gradually liberalized. In particular, the central government of China officially initiated a so-called “adjustment and reform” policy in 1979, in order to promote quicker growth in the industrial and agricultural sectors. It was under this background that the old system of administered prices was increasingly liberalized. Consequently, prices of both agricultural products and industrial products increased considerably in the early 1980s, which inevitably passed through the production chain and generated high inflation.

In conjunction with growing inflation, the growth rate of real output in the early 1980s was also accelerating. Countercyclical macro policies, however, were not implemented in a timely and effective manner. Despite a transitory drop in 1982, growth rates of both M2 and domestic credit in China exhibited an upward trend in late 1982, with unprecedented accommodative levels of 40% in 1985 and nearly 50% in 1986. As a result, there was evidence of overheating, with inflation peaking in 1985–1986. The tightening of credit controls in 1986 dampened inflation, but it was effective for only a very short period of time. Due to further liberalization and deregulation of prices in 1987, inflation rebounded to a high of 25% in 1988. In response to the extraordinary inflation, the central government tightened money and credit supply and reduced fixed investment substantially. These tighter monetary policy conditions towards the end of the 1980s successfully curbed inflation.

Although the tightening macro policy in the late 1980s had the effect of cooling down inflation (and economic growth), it proved to be too constrictive. Due to the strict credit control in 1988 and 1989,
the industrial sector witnessed a substantial reduction in its output in the ensuing three years, which consequently caused a serious liquidity problem among enterprises in China. As a result, both economic growth and inflation declined to a relatively low level (below 5%) in 1991.

In the early 1990s, however, agricultural prices were adjusted upwards to market levels and price controls were eliminated in industrial and retail sectors (Narayan et al., 2009). In the meantime, the central government encouraged investments by loosening credit control aggressively, with money supply growing at over 50%. The proactive policy led inflation to increase in 1992 and peak in 1994 (as attested in Fig. 1). Following a number of tightening policy measures in 1994, inflation started to decelerate in 1995 and continued to decline until the late 1990s.

Since the end of the 1990s, China has experienced two periods of mild deflation occurring from 1998 to 2000 and in 2009, with relatively low and stable inflation in the rest of the years of the new century. It may be worth noting, however, that the GDP in 1997 was relatively low and stable in 2001. In practice, the GDP deflator is 1.2% in 1991.

The foregoing description indicates that the evolution of inflation in China moved in conjunction with monetary policy changes (as well as supply and demand shocks) over time. Since China has witnessed profound changes in its monetary policy regime over the past three decades, it is important to examine the link between the instability of the inflation process and monetary policy shifts. The following section therefore, embarks on structural break tests.

3. Inflation persistence and structural break tests

3.1. Univariate regression

We examine structural changes in China’s inflation process using a battery of break tests, applied to simple autoregressive (AR) models. The tests look for structural changes in the coefficients in the AR model

$$\pi_t = c + \alpha(L)\pi_{t-1} + \varepsilon_t, \quad \pi_t = \left\{ \begin{array}{ll} \varepsilon_t + \alpha_1(L), & t \leq \tau \\ \varepsilon_t + \alpha_2(L), & t > \tau \end{array} \right. \quad (1)$$

where $c$ denotes an intercept, $\alpha(L)$ is the polynomial in lag operator, $\pi_t$ is the rate of inflation, and $\varepsilon_t$ is a serially uncorrelated error term. Note that $\tau$ indicates the break date in the conditional mean of the underlying inflation process.

The representation in Eq. (1) summarizes the impact of past levels of inflation on current inflation through the coefficient $\alpha(1) = \alpha_1 + \alpha_2 + \ldots + \alpha_p$, where $p$ is the optimal lag order based on AIC and serial correlation test. By construction, parameter $\alpha(1)$ measures the level of inflation persistence, that is, how long it will take inflation to return to its steady state level after being disturbed. Because the persistence parameter effectively determines the nature of the inflation process, its structural stability is our focus.

In a traditional Chow test, one has to set a specific break point based on a priori knowledge about the potential break date. In our analysis, however, we do not assume any prior knowledge about potential break dates. We employ the heteroskedasticity-robust Quandt–Andrews unknown break point test, originally introduced by Quandt (1960) and later developed by Andrews (1993) and Andrews and Ploberger (1994). The test statistics of the Quandt–Andrews family include Supreme likelihood ratio statistic (Sup-LR), exponential LR (Exp-LR), and Average LR (Ave-LR). All three tests are Andrews family are summarized in Table 1, which reports $p$-values of these tests are computed using Hansen’s (1997) method.

By construction, the Sup-LR statistic is the maximum LR-statistic for testing a break through all possible break points over the specified search range ($\tau$), which is given by

$$\text{SupLR} = \sup_{\tau \in [\tau_{\min}, \tau_{\max}]} \text{LR}(\tau) \quad (2)$$

where $\text{LR}(\tau)$ denotes the sequential LR-statistics testing for the null hypothesis of no structural break in the underlying parameter. We set a search interval $\tau = [0.15, 0.75]$ for the full sample $T$ to allow a minimum of 15% of effective observations contained in both pre- and post-break periods. The Exp-LR and Ave-LR statistics are then computed as the average and exponential of the corresponding sequential LR-statistics.

The results of the structural break tests of the Quandt–Andrews family are summarized in Table 1, which reports $p$-values of the test statistics. The tests are performed on all of the coefficients (overall),

**Fig. 1. GDP deflator inflation in China: 1979Q1–2009Q4.** Source: The author’s calculations.
the intercept, and the persistence coefficient. Specifically, the first
statistic tests for the stability of all coefficients in Model (1). The
second statistic tests for stability of the intercept term assuming the
remaining coefficients are constant. Likewise, the third statistic tests
for constancy of the persistence parameter assuming the other
coefficients are constant.

The structural break tests for the overall parameters in Table 1
indicate statistically insignificant evidence of a structural break in
the inflation process. However, the corresponding break date (2003Q4)
ocurs at an extreme of the search interval, indicating that the
insignificant result of the associated break tests is probably distorted
by the extreme sample problem. Indeed, investigating the other two
tests in the regressions reveals that there is a significant structural break
in the persistence parameter (and the intercept) around 1995Q4.

Overall, the Quandt–Andrews statistics in Table 1 suggest that the AR
process of the GDP deflator inflation is unstable over 1979–2009. In particular,
we find that the structural break in the inflation process
concentrates in the persistence parameter (and the intercept as well) in
late 1995; and so to assess the changing nature of inflation persistence,
we investigate the estimates of inflation persistence over the two
different periods based on the identified structural break date.

However, the estimation of the persistence parameter in the AR
model is complicated by the skewing distribution bias problem as
shown in Hansen (1999). In effect, the standard asymptotic distribu-
tions for OLS coefficients and t-statistics in AR models become system-
atically poorer approximations to their true finite-sample distribu-
tions as the true value of the persistence parameter increases.
In particular, point estimates become increasingly downward-biased
and their distribution becomes more skewed to the left as the true value
of the persistence parameter moves closer to unity (O’Reilly and
Whelan, 2005).

To rectify this problem, we also use Hansen’s (1999) grid-bootstrap
method to calculate the bootstrap distribution over a grid of values of
the autoregressive parameter in AR models and then obtain unbiased
median estimates for the inflation persistence parameter. The estimation
results for inflation persistence over different samples are reported
in Table 2, from which it is evident that the persistence parameter for the
GDP deflator inflation before and after 1995 indeed differs markedly.
According to the grid-bootstrap median unbiased estimates, inflation
persistence was high (0.905) before the structural break, and comparatively
low (0.784) after the break point. The OLS estimates of the persistence parameter provide a more striking contrast between the
two different samples (0.853 versus 0.660).

3.2. Conventional Phillips curve regression

Up to this point, we have been concentrating on estimating the
persistence parameter in the univariate inflation process, but within
this focus we may have effectively squeezed a potential real driving

variable into the error term. The exclusion of any real driving variables
may result in spurious findings with a high value for the inflation
persistence parameter owing to omission of possible feedback from the
real variable to inflation. In addition, recent research (e.g., O’Reilly
and Whelan, 2005)) shows that unless we condition 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 the real output gap on inflation. To
achieve this, we augment Model (1) by adding a measure of the
output gap, which we have constructed using a Hodrick-Prescott
filter. That is, we estimate inflation persistence in a conventional
Phillips curve specification in the spirit of Stock and Watson (1999), viz.

\[ \pi_t = \frac{c}{\gamma} + \alpha (\Delta \pi_{t-1} + \beta y_{t-1} + u_t) \]

where \( y_t \) denotes the output gap and \( u_t \) denote a random error.

Based on the specification of Model (3), we first perform the
Quandt–Andrews structural break tests and the results are consistent
with Model (1). In particular, the persistence parameter for the GDP
deflator inflation in the Phillips curve specification also manifests
a significant structural change (the corresponding p-values of the
structural break tests are smaller than 5%). We then examine
subsample estimation results of the Phillips curve specification for
the underlying inflation series, using both the grid-bootstrap and the
OLS estimators. It turns out, though, that the overall pattern of our
results concerning the persistence parameter is little changed by
inclusion of the output gap. The result worth noting is that the
reduction of inflation persistence in Model (1) is more pronounced
than that shown in Table 2, using either the grid-bootstrap method or
the traditional OLS estimator. Taken as whole, the general pattern of
structural changes in inflation persistence in the Phillips curve
specification does not alter that in the AR process of Model (1),
suggesting the robustness of our finding.

3.3. Robustness assessments

The robustness of the structural change in inflation persistence is
checked in two ways, firstly by using non-spliced data over 1992–
2009, and secondly by applying a refined (repartition) procedure for
multiple structural break tests. These are elaborated upon below.

First, inflation data prior to 1992 are constructed from annual
estimates and there is a potential concern that our revealed break
result picks up this change in the spliced data and is not a structural
development. To check whether the break point in 1995 is influenced
by the data splice issue, we use non-spliced data over 1992–2009
and perform the Quandt–Andrews unknown break point test for the AR
model over this sample period.

The relevant results are reported in Table 3, which show that
exactly the same break in the persistence parameter is detected. In
addition, the recent study by Zhang and Clovis (2009b) also found a
break point in China’s inflation persistence (consumer price and retail
price inflation) around the same time, suggesting the timing of the
discovered break is not influenced by the splice in the GDP deflator
data.

Our second robustness assessment is to check whether there are
multiple structural breaks in inflation dynamics. We sequentially

5 The output gap is calculated as 100 × [ln(GDP) − ln(Trend)] where Trend is the
Hodrick-Prescott smoothed series (with penalty parameter 1600). In practice, the
output gap is treated as a fixed regressor in the grid-bootstrap estimation. Detailed
results for these tests are not reported for brevity, but they are available upon request.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Grid-bootstrap</th>
<th>OLS</th>
</tr>
</thead>
<tbody>
<tr>
<td>Full sample</td>
<td>0.877</td>
<td>0.844</td>
</tr>
<tr>
<td>Pre-break</td>
<td>0.565</td>
<td>0.853</td>
</tr>
<tr>
<td>Post-break</td>
<td>0.784</td>
<td>0.660</td>
</tr>
</tbody>
</table>

Notes: *, **, and *** indicate statistical significance at 10%, 5%, and 1% levels, respectively.
apply the Quandt–Andrews unknown break point test to subsamples and perform the refined (repartition) procedure suggested by Bai and Perron (1998). That is, if the null hypothesis of stability is rejected, then a break point is estimated as the date corresponding to the Sup-LR statistic and the sample is divided into two subsamples on this date. The stability tests are then applied to each subsample to estimate any additional breaks. If multiple structural break points are indicated, their dates are refined (re-estimated). The resulting estimate has the same convergence rate as in the case of simultaneous estimation of multiple breaks. The corresponding results (not reported here for brevity), however, suggest no further significant structural break in the underlying model of inflation dynamics.

4. Inflation expectations

The empirical results in the preceding section indicate that the persistence of the GDP deflator inflation experienced a significant reduction around 1995, which is likely to be associated with some policy regime shifts in China. In particular, the distinct structural change in the persistence parameter in 1995 coincides and is consistent with the strengthening role of the People’s Bank of China (PBOC, the central bank of China) in implementing monetary policy in the mid-1990s. Historically, the PBOC did not begin functioning as a central bank until 1978 due to the destructive Cultural Revolution (1966–1976). Although it assumed its responsibility as a central bank in 1983, its status and authority as a modern central bank was not legally confirmed until the Law of the People’s Republic of China on the People’s Bank of China was enacted in March 1995.

Since then, there have been significant improvements in China’s monetary policy. For instance, the establishment of a unified interbank money market in 1996 has facilitated liquidity adjustment for the PBOC. In 1997, China announced “The Regulations on the Monetary Policy Committee of the People’s Bank of China,” which clarified the rights and responsibilities of the monetary policy committee in the central bank. In 1998, the PBOC replaced quota management of credit with assets-to-liabilities ratio management. Open market operations were also resumed in 1998. Hence, the PBOC can adjust its intermediate target by issuing central bank bills and using repurchasing agreements to make collateralized loans to primary dealers.

In January 1999, the central bank of China abolished its branches at provincial and municipal levels and set up nine regional branches to primary dealers.

In this setup, the trend component $\tau_t$ follows a random walk (as indicated in the second equation of (4), so that a stochastic shock to this component will persist permanently and hence affect the inflation trend going forward. The transitory component $\eta_t$ is serially uncorrelated, indicating that such shocks are temporary and lead to only transitory fluctuations around the long-run trend. Note that Model (4) also allows the volatility of the two kinds of shocks to vary over time, as indicated by the last two equations in (4). By definition, the trend component represents a long-run outlook of inflation and thus can be used to mimic inflation expectations.

The estimation results of Model (4) are illustrated in Figs. 2 and 3. Fig. 2 compares the dynamic evolution of stochastic volatility of the trend and transitory components, while Fig. 3 shows the estimated trend component of the GDP deflator inflation and the actual inflation series. As attested by Fig. 2, the importance of the trend shocks, relative to that of the temporary shocks, started to rise at the end of the 1970s, declined in the late 1980s, grew in the first half of the 1990s, peaked in the middle of the 1990s, declined again in the next ten years, and finally, just prior to the recent crisis episodes it exhibited a short rising period to 2007.

When the relative importance of the trend shocks became high, as they did in the early 1980s and the mid 1990s, inflation became highly persistent. Under such conditions, if inflation goes up, the trend component will rise in tandem so that inflation will remain high. In contrast, when the trend shocks were relatively less important, as was the case after the end-1990s, a change in inflation tended to diminish more quickly, indicating lower inflation persistence.

However, it should be noted that although the importance of the trend component appears to fall in the new century, the estimates suggest its communication with the public and improves its transparency. Once the central bank is more transparent, the public will be more confident about the central bank’s ability to control inflation. As a result, market expectations of future inflation will be better anchored.

To verify the above conjecture, we carry out two exercises. The first exercise is to investigate whether the transparency of the PBOC has improved. To this end, we follow Eijffinger and Geraats (2006) and calculate a transparency index for the PBOC over pre- and post-1995 periods respectively. The results are reported in Table 4 (where higher values represent greater clarity), which shows that the total transparency index is only 2.5 prior to 1995, while it increases to 12.5 after the structural break date.

Our second exercise is to estimate the time series of long-run inflation expectations. We employ Stock and Watson’s (2007) unobserved components model with stochastic volatility (UCSV). This model utilizes an intriguing method in decomposing an inflation series into one long-run trend and one transitory component. The underlying model can be summarized as follows:

\[
\begin{align*}
\pi_t &= \tau_t + \eta_t, \quad \text{where} \quad \eta_t = \alpha_{\eta} \xi_{\eta,t} \\
\tau_t &= \tau_{t-1} + \epsilon_t, \quad \text{where} \quad \epsilon_t = \alpha_{\epsilon} \xi_{\epsilon,t} \\
\ln \sigma_{\eta,t}^2 &= \ln \sigma_{\eta,t-1}^2 + \omega_{\eta,t} \\
\ln \sigma_{\epsilon,t}^2 &= \ln \sigma_{\epsilon,t-1}^2 + \omega_{\epsilon,t}
\end{align*}
\]

This formulation allows for the trend and transitory components to have different volatilities and for the trend component to be serially uncorrelated, which is consistent with the notion that the trend component represents a long-run outlook of inflation and thus can be used to mimic inflation expectations.

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<table>
<thead>
<tr>
<th>Table 3</th>
<th>Results ($p$-values) of Quandt–Andrews unknown break point test over 1992–2009: non-spliced data.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Sup-LR</td>
<td>0.456</td>
</tr>
<tr>
<td>Exp-LR</td>
<td>0.219</td>
</tr>
<tr>
<td>Ave-LR</td>
<td>0.138</td>
</tr>
<tr>
<td>Break date</td>
<td>1995Q3</td>
</tr>
</tbody>
</table>

Note: Notations follow Table 1.

<table>
<thead>
<tr>
<th>Table 4</th>
<th>Transparency index of the PBOC.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Before 1995</td>
<td>1</td>
</tr>
<tr>
<td>After 1995</td>
<td>3</td>
</tr>
</tbody>
</table>

Notes: The index is calculated based on Eijffinger and Geraats (2006).
that it was still large enough to be economically meaningful. As we can see from Fig. 3, the estimated trend of the GDP deflator inflation had drifted up since early 2007, implying a potential rise of inflation in coming periods. Although such estimates are inevitably imprecise to some extent (Mishkin, 2007), our results suggest that the structural change in inflation persistence is connected with changing inflation expectations, while the long-run trend of the GDP inflation series may not yet be perfectly stable.

5. Counterfactual simulations

The foregoing section describes important changes in the conduct of China’s monetary policy and the resultant changes in inflation expectations. We also provide preliminary evidence about the link between the changes in monetary policy (and the resultant changes in inflation expectations) and inflation persistence. These descriptions suggest that the change in the observed significant reduction in inflation persistence may reflect a fundamental shift in monetary policy whereby the PBOC now acts systematically to stabilize inflation around a potential long-run target and has gained credibility with the public in believing that it will continue to do so into the future. This causal relationship between the change in systematic policy and the structural shift in inflation dynamics reflects the essence of the Lucas critique.

The same interpretation for the change in inflation persistence is proposed by Taylor (2000) and Williams (2006) for the US economy. An alternative interpretation found in the analysis of Stock and Watson (2007) is that the economy has simply experienced a run of good luck and the decline of inflation persistence may only reflect the fact that the nature of “shocks” has changed.

These different interpretations have distinct implications for policy analysis. If the structural change in inflation dynamics is attributed to factors other than monetary policy, it implies that the PBOC could respond less to shocks and continue to be confident that inflation would remain at a low level. However, if the change in inflation dynamics is attributed to improved monetary policy intervention by the PBOC, then Stock and Watson’s policy conclusion is unwarranted. On the contrary, it implies that better policy has resulted in better anchoring of inflation expectations because of which inflation has become less persistent and is now less responsive to shocks.

To explore these alternative interpretations of the changing inflation dynamics, we construct a Vector Autoregression (VAR) model and conduct counterfactual simulations to quantify contributions of monetary policy and random shocks to the observed structural change in inflation persistence. We note that although the PBOC has recently promoted development of market-based interest rates as policy instruments (e.g., Shanghai interbank offered rate was launched by the PBOC in January 2007), quantity-based monetary instruments remain the main instruments of the PBOC, as explicitly stated in the Monetary Policy Report published quarterly by the PBOC and shown in Burdekin and Siklos (2008) and Geiger (2008). Therefore, our baseline VAR model incorporates quarterly data for the growth rate of real GDP, inflation, and the growth rate of M2 to
capture the dynamics of real economic growth, inflation, and monetary policy.\textsuperscript{6} By construction, the equation with the growth rate of M2 as the dependent variable in the VAR system is used to capture the dynamic reactions of monetary policy to inflation and to real economic growth.

We impose the identified break point (i.e. 1995Q4) as a break date and use the VARs estimation over 1979–1995 and 1996–2009 to simulate the series of the GDP deflator inflation for the respective subsamples. Through this exercise, we attempt to assess whether changes in propagation of the system (VAR parameters) or the impulse (the innovation covariance matrix) can account for the observed reduction in inflation persistence.

Specifically, each VAR is formed as follows:

\[ Y_t = \Phi(L)Y_{t-1} + \varepsilon_t, \quad \text{var}(\varepsilon_t) = \Sigma \]

where \( Y_t \) is a vector time series and the subscript \( i = 1, 2 \) denotes the first and second subsample (the intercept is omitted in the equation for notational convenience but is included in the estimation); \( \Phi(L) \) denotes the polynomial of the lag operator with optimal lag order determined by AIC. The estimates of the parameters and the innovation variance matrix in (5) are then used to generate inflation series over different sample periods. Note that the simulation process assumes the structural innovations (denoted \( e \)) implied by the reduced VAR model follow vector Gaussian white noise which is linked to the shocks in (5) through \( \Sigma = A_0^{-1} \Sigma_0(A_0^{-1})' \), where \( A_0 \) is the implicit contemporaneous parameters vector in the counterpart structural VAR model to (5) and is recovered by using the standard Cholesky decomposition method.

The simulated inflation series are then used to estimate the persistence parameters in the univariate model. By using different \( \Phi \) and \( \Sigma \), it is possible to compute counterfactual persistence of inflation that would have arisen had either \( \Phi \) or \( \Sigma \) taken different values. For example, \((\hat{\Phi}_1, \hat{\Sigma}_1)\) is the estimated inflation persistence in period one, and \((\hat{\Phi}_2, \hat{\Sigma}_2)\) is the inflation persistence that would have occurred had the lag dynamics been those of the second period and the error variance matrix been that of the first period. Other cases are defined analogously.

Table 5 summarizes the VAR-based estimates of inflation persistence for the four possible permutations of the estimated VAR coefficients and variance matrices. Columns labeled \((\hat{\Phi}_1, \hat{\Sigma}_1)\) and \((\hat{\Phi}_2, \hat{\Sigma}_2)\) respectively contain the VAR-based estimates of the first and second-period sample inflation persistence which are quite close to the respective sample estimates also shown in Table 5. The columns labeled \((\hat{\Phi}_1, \hat{\Sigma}_2)\) and \((\hat{\Phi}_2, \hat{\Sigma}_1)\) contain the counterfactual estimates. Taking the OLS estimates for example, the counterfactual combination of first-period dynamics and second-period shocks \((\hat{\Phi}_1, \hat{\Sigma}_2)\) produces an estimated persistence of 0.841, which is close to the first-period persistence estimate (0.889). In contrast, the second-period lag dynamics and first-period shocks produce an estimated inflation persistence of 0.702 which is much less than the first-period persistence estimate.

That said, had the lag dynamics of the post-1995 period occurred in the pre-1995 period, the GDP deflator inflation would have been much less persistent. Our empirical results indicate that the changes in the parameters of the VAR system—the dynamic propagation of the system—can account for most of the observed reduction in inflation persistence, while the contribution from random shocks is much smaller. The results of the grid-bootstrap estimation in Table 5 suggest the same conclusion.

Overall, the counterfactual analysis indicates that improved monetary policy is the main contributor to the reduced inflation persistence in China. According to Mishkin (2007), better monetary policy reduces inflation persistence through the anchoring of inflation expectations. In this aspect, monetary policy is the source of the change in the evolution of long-run inflation expectations, per se. This explanation is consistent with the stylized facts in China. As Sections 2 and 4 explain, during the 1980s and 1990s, China maintained a policy stance that was possibly too easy and allowed inflation expectations to drift markedly up over many years. Since the late 1990s, however, the PBOC has increased its commitment to price stability in both words and actions. The preemptive strikes against fluctuations in inflation by the PBOC since the late 1990s have not only kept inflation low and stable, but have also anchored long-run inflation expectations, which effectively helped reduce inflation persistence.

6 Using output gaps as alternative measures for the real variable in the VAR model does not alter the finding.

### 6. Conclusions

In this paper, we have presented evidence of structural instability over time in dynamic equations for inflation in China. The significant shift in the inflation persistence parameter is shown to be the response of inflation expectations to the sequence of reforms in China’s monetary policy regime that have taken place since the mid 1990s. This finding suggests that the Lucas critique is not a purely theoretical result, but rather a warning that underscores the importance of applying appropriate structural stability tests to dynamic models for inflation in China.

The baseline finding in the present paper also implies that inflation in China is now less persistent and less responsive to inflationary shocks. However, since the structural change in the inflation process is attributed mainly to better monetary policy and the associated better inflation expectations, it is also possible for high inflation to strike back at China in the absence of a determined effort by the monetary authorities to continue to manage inflation expectations. The estimation results of the UCSV model for the underlying inflation series show that the importance of the trend component is still large enough to be economically meaningful, indicating a potential rise of inflation in coming periods. Therefore, systematic monetary policy improvements for managing market expectations of future inflation and hence short-run inflation dynamics are clearly still warranted in China.

It should be noted that our search for the causes of the significant structural break in Chinese inflation dynamics is not intended to be exhaustive. Other factors may also influence inflation, and some of these may provide other possible explanations for the recent change in Chinese inflation dynamics. For example, increased globalization and competition may have lowered the sensitivity of domestic inflation to alternative shocks. Another factor possibly influencing inflation in China around the mid 1990s is changing exchange rate regime. Just prior to the identified structural break point in 1995, there was a large depreciation of the RMB against the US dollar, which marked a distinct shift in exchange rate regime in China. Therefore, it could be fruitful for future research to adopt a more structural framework incorporating, when tractability allows, all relevant factors pertaining to inflation dynamics in China. Studies toward this direction may provide more compelling results that may complement the present research.

### Table 5

<table>
<thead>
<tr>
<th>Method</th>
<th>((\hat{\Phi}_1, \hat{\Sigma}_1))</th>
<th>((\hat{\Phi}_2, \hat{\Sigma}_2))</th>
<th>((\hat{\Phi}_1, \hat{\Sigma}_2))</th>
<th>((\hat{\Phi}_2, \hat{\Sigma}_1))</th>
</tr>
</thead>
<tbody>
<tr>
<td>OLS</td>
<td>0.889</td>
<td>0.510</td>
<td>0.841</td>
<td>0.702</td>
</tr>
<tr>
<td>Grid-bootstrap</td>
<td>0.937</td>
<td>0.558</td>
<td>0.881</td>
<td>0.753</td>
</tr>
</tbody>
</table>

Notes: The table reports point estimates for each permutation of the counterfactual analysis. Initial values are actual observations; the first 100 obs. of the simulated series are discarded to avoid sensitivity; simulated sample sizes are consistent with the respective subsamples.
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References


