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Deflation and Relative Prices: Evidence from Japan and Hong Kong

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Abstract

We test the menu cost model of Ball and Mankiw (1994, 1995) on data from the inflation and deflation periods in Japan and Hong Kong. We calculate the moments of the distribution of price changes using a random split procedure to overcome the bias noted by Cecchetti and Bryan (1999). The key implication of the model, that the parameter on skewness is positive during both inflation and deflation, is strongly supported. The data are less clear on, but does not reject, the hypothesis that the parameter on the standard deviation changes sign between inflation and deflation periods.

Keywords: inflation, deflation, menu costs, Hong Kong, Japan

JEL Numbers: E31

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1. Introduction

A large body of empirical evidence indicates that inflation is positively correlated with the cross-sectional variability of relative prices changes.¹ While this relationship was already noted by Frederick C. Mills (1927) in US data, the newer literature was started by Herbert Glejser (1965), who demonstrated that there was a positive relationship between the standard deviation of relative prices and inflation using three different samples of data. Further evidence in support of this effect was established by Daniel R. Vining and Thomas Elwertowski (1976), who studied annual data for wholesale and retail prices in the United States between 1948 and 1974.

The positive association between inflation and relative price changes is difficult to explain with a classical view postulating price flexibility, which sees relative price changes as solely determined by real factors and the overall inflation rate as determined by excessive money growth. To generate a positive relationship between inflation and relative price variability, some additional element is needed. In a widely cited paper, Stanley Fisher (1981) discusses a number of potential explanations for the observed correlation. Two of these have attracted considerable attention in the literature on inflation and price dispersion.²

The first is based on the imperfect information model of Robert E. Lucas (1973). In this framework, agents are distributed across “islands.” They observe only local prices and face the problem of distinguishing between local and aggregate shocks. Since they are unable to do so perfectly, unobserved aggregate shocks are partially misinterpreted as local shocks and are therefore associated with supply responses and relative price changes.

The second explanation, which is increasingly dominant in the literature and the focus of this paper, is based on the assumption that firms face fixed costs of changing prices (“menu costs”), which lead to infrequent price adjustments. Given these costs, an increase in the inflation rate leads to more dispersion of relative prices as only some firms’ change prices. One important consequence of the fact that inflation drives relative prices apart is that

¹ A related finding is that the level and variability of inflation are correlated, which has also spurred a large literature. See for instance Arthur M. Okun (1971) and Dennis E. Logue and Thomas D. Willet (1976) for early studies.

² Fischer (1981) considers several other potential explanations for observed relationship, including downward price inflexibility and the possibility that non-monetary shocks lead, potentially through policy responses, to more inflation and stronger relative price variability as different industries adjust at different speeds.

monetary policy makers should stabilize the aggregate price level to maximize agents' welfare, as argued by Michael Woodford (2003).

Both these strands of the literature attribute variations over time in price dispersion to movements in inflation, and empirical work typically proceeds by regressing the degree of price dispersion on inflation and other variables. In the more recent literature on menu costs, Laurence Ball and N. Gregory Mankiw (1994 and 1995) have argued that the association between inflation and price dispersion arises because firms adjust prices only in response to large shocks to their desired relative price in the presence of menu costs. Importantly, this leads to a relationship in which both the standard deviation and skewness of individual price changes determine the mean of the distribution of price changes.³

To see this, suppose that the distribution of price changes is symmetric and that the trend rate of inflation is zero. In this case, the number of firms that would like to raise prices is equal to the number of firms that would like to cut prices, so that the overall level of prices remains stable irrespective of the importance of menu costs and the variance of relative prices. If instead the distribution of relative prices changes is skewed to the right, there are few firms that would like to increase their prices a lot and many firms that would like to cut them a little. In the presence of menu costs, the firms that would like to raise prices a lot will choose to do so, while the firms that would like to cut prices a little will refrain from doing so. As a consequence, a positive skewness will be associated with rising, and a negative skewness with declining, prices.

While in the above thought experiment only skewness matters for inflation, the dispersion of prices is important for two reasons. First, as argued by Ball and Mankiw (1995), while the variance of individual prices may not have an independent effect on the rate of change of prices, it reinforces the effect coming from skewness. Suppose that the distribution is skewed to the right and the variance of individual prices increases. If so, there will be a greater number of firms that would like to raise prices and a greater number that would like to cut them. However, since the former are more likely than the latter to be willing to incur the menu costs, it follows that an increase in price dispersion is associated with a stronger tendency for prices to rise. Similarly, if the distribution is skewed to the left, an increase in price dispersion

³ Interestingly, David R. Vining, Jr. and Thomas C Elwertowski (1976, p. 703) noted already years before this that empirically there is a positive relationship between the skewness and the mean of the distribution of price changes.

will raise the likelihood that average prices are falling. This suggests that skewness interacted with degree of price dispersion may be important determinant of inflation.

Second, the dispersion of price changes can also be important if the trend rate of inflation is positive even if the distribution of price changes is symmetric. Ball and Mankiw (1994 and 1995) show that in this case an increase in the dispersion of prices raises the rate of inflation. The reason for this is that firms whose equilibrium relative price increases will not change nominal prices but merely wait and let inflation erode their relative price gradually. By contrast, firms who have experienced a negative shock to their equilibrium prices can not rely on this adjustment mechanism and will instead raise their prices. Thus, in the presence of trend inflation, an increase in the variance of relative prices will raise inflation. Of course, the argument is symmetric: under trend deflation, an increase in the variance of relative prices will increase the rate of deflation.

It is consequently possible to test the menu cost model by exploring whether the sign of the impact of the dispersion of relative prices is positive in sample periods of inflation and negative in periods of deflation. Apparently, this has not been done in the literature, which motivates the present paper.

In sum, the menu cost model implies that there should be a positive relationship between the degree of asymmetry in the distribution of price changes, as captured by the skewness, irrespectively of the behaviour of trend inflation. Furthermore, holding the skewness constant, increases in the degree of price dispersion will tend to amplify the underlying rate of change of prices. Thus, an increase in price dispersion will raise inflation in periods with inflation, raise deflation in periods of deflation, but have little effect in situations in which the economy is at price stability.

Note furthermore that while the skewness plays an important role in the Ball-Mankiw model, it does not do so in the Lucas model. It is consequently possible to distinguish between these models empirically by focussing on whether the skewness matters for inflation.

This paper expands the existing literature in two ways. First and as noted above, it is of interest to study the relationship between the rate of change of prices, price dispersion and skewness using data from episodes in which the economy switched between inflation and deflation. The reason for this is that the sign of the parameter on price dispersion in an inflation equation should depend on whether the economy is in inflation or deflation. This hypothesis does not appear to have been tested in the literature. In this paper we do so using

disaggregated monthly CPI data covering the last two decades for Japan and Hong Kong, which experienced inflation until the middle of 1998 and then underwent a long period of protracted deflation that ended between 2004 and 2005.

Second, we propose a method to overcome the problem identified by Cecchetti and Bryan (1999), who show that regression results for the inflation rate and higher-order sample moments can be subject to severe bias in finite cross-sectional samples. In particular, high kurtosis of the relative-price change distribution may lead to a spurious relationship between the inflation mean and its skewness. We use a random cross-section sample split for calculating the mean and the higher-order moments of the distribution of price changes to mitigate this bias. Furthermore, we show that the bias problem is mitigated if the variance of the long-run or trend inflation rate is high, which is an additional reason for why the use of inflation and deflation period data (which contain considerable long-run variation of the inflation rate) is interesting from a statistical point of view.

The paper is organized as follows. In the second section we provide an overview of inflation developments in Japan and Hong Kong from the 1980s onward. In Section 3 we review and extend the work of Bryan and Cecchetti (1999), who argue that the parameters in Ball-Mankiw regressions of inflation on the skewness and standard deviation of the cross-sectional distribution of price changes can be subject to severe bias.

Section 4 contains the empirical results we obtained for Japan and Hong Kong using monthly data from the early eighties to the recent past. We find that the bias problem is important as OLS estimates are at times strikingly different from those obtained using our alternative procedure, which yields a positive and highly significant coefficient on skewness in both inflation and deflation periods and in both countries. These results provide strong support for the menu cost hypothesis but are difficult to rationalise under the Lucas model.

We therefore show that the point estimate of the coefficient on the standard deviation of price changes sign from positive to negative as Hong Kong falls into deflation, as suggested by the menu cost model. In Japan, by contrast, the point estimate is positive in both periods. However, the estimated parameters are all insignificant. It may be due to the short sample period, coupled with the fact that, at least in Japan, the change in the behaviour of prices between the inflation and deflation period was not so marked (as we discuss below).

Finally Section 5 concludes.

2. Deflation in Japan and Hong Kong

As a backdrop to the analysis below, we review next price developments in Japan and Hong Kong before and during the deflation period.

Figure 1, which shows the rate of change of consumer prices (measured over four quarters) in the two economies, warrants two comments.⁴ First, inflation in the two economies followed similar time paths, although the amplitude of the fluctuations was much less pronounced in Japan than in Hong Kong. Thus, inflation fell from the beginning of the sample toward the mid-1980s and then peaked in the early 1990s. It subsequently declined before turning to deflation around mid-1998 in both economies.⁵ The deflation ended in the middle of 2004 in Hong Kong and at the end of 2005 in Japan.

Second, Hong Kong had a much more severe brush, although somewhat briefer, with deflation than Japan, which arguably is best described as being close to absolute price stability throughout this period. The monthly statistics reveal that prices in Hong Kong fell a cumulative 16.3% between the peak of the CPI in May 1998 and the trough in August 2003, with the peak rate of deflation being 6.3%. The consumer price index (excluding fresh food) in Japan fell by 4.9% between September 1997 and February 2005, with a peak rate of deflation of 1.5%. Measured by the change over twelve months, Hong Kong experienced 68 months, and Japan 88 months, of deflation.

Overall, the two deflation episodes share many similarities, in particular, they were protracted, lasting between six to eight years. Thus, while the onset of deflation might have been unexpected, over time expectations of further price declines took hold.

--- Figure 1 here ---

3. Methodological issues

The data analysed below consist of N time series with T observations on the components of the CPI. We denote the logarithm of series i at time t by p_{it} . In what follows we let μ_{jt} and m_{jt} denote the true and estimated values of the j th moment at time t . The economy-wide inflation rate can then be defined as:

⁴ For Japan we use the CPI excluding fresh food; for Hong Kong we use the Composite Consumer Price Index (CCPI).

⁵ Japan also experienced declining prices between the middle of 1995 and the end of 1996.

$$\pi_t = m_{1t} = \sum_{i=1}^N w_i \Delta p_{it} \quad (1)$$

where w_i denotes the weight of component i in the CPI. In addition to this first moment of the price-change distribution we define centred and weighted higher order moments as:

$$m_{rt} = \sum_{i=1}^N w_i (\Delta p_{it} - \pi_t)^r, \quad r = 2, 3, \dots \quad (2)$$

The standard deviation (STD) as well as the coefficients of skewness (SKEW) and kurtosis (KURT) are obtained by dividing the second, third and fourth moment by $(m_{2t})^{r/2}$. Using CPI-weights instead of equal weights is appropriate since it takes into account the relative importance of price changes in the subcategories of the CPI.

Bryan and Cecchetti (1999) showed that spurious relationships between the mean and higher moments may arise in small cross-section samples and demonstrated that the highly significant correlation between inflation and the third moment found in US CPI and PPI data is caused by this problem. This suggests that the conclusion by Ball and Mankiw (1995) that the menu-cost model explains the behaviour of aggregate US inflation may be incorrect. We next review the arguments of Bryan and Cecchetti, discuss their consequences for time-series regressions of inflation on the sample moments of the distribution of relative-price changes and introduce an IV estimator to deal with the potential bias.

To illustrate the problem, consider a panel with cross-section dimension N and time dimension T . We assume that the data, denoted by x_{it} (in the case above, we have that $x_{it} \equiv \Delta p_{it}$), is driven solely by one common element, namely a time-varying mean Z (that is, $\sigma_Z^2 > 0$). Formally:

$$\begin{aligned} Z_t &= E_i x_{it} \\ E(Z_t) &= 0 \\ E(Z_t)^2 &= \sigma_Z^2, \\ E(Z_t Z_{t-s}) &= 0, \quad s \neq 0 \end{aligned} \quad (3)$$

The x_{it} are identically, independently and symmetrically distributed across i and stationary over t . The assumptions of symmetry and zero unconditional mean of Z are only made to simplify the exposition and could be easily replaced by a non-zero mean and a skewed distribution. Moreover, we assume that the x_{it} have higher-order population moments around

the mean denoted by μ_r . The time series of sample moments denoted by m_{rt} are obtained in analogy to equations (1) and (2) with equal weights $1/N$ as we have a pure random sample.

In this simple framework there is by construction no relationship between the mean and the higher-order moments of the cross-section distribution over time. Thus, any correlation between the estimates of the mean and higher moments is solely due to estimation error. Bryan and Cecchetti (1999) consider the correlation of the mean and the third moment which is T -asymptotically:

$$\rho_{1,3} = \frac{E(m_{1t}m_{3t})}{(E(m_{1t})^2 E(m_{3t})^2)^{1/2}} \quad (4)$$

For the i.i.d. case the three relevant expected values are:⁶

$$E(m_{1t}m_{3t}) = \left[\frac{1}{N} - \frac{3}{N^2} + \frac{2}{N^3} \right] \mu_4 \quad (5)$$

$$E(m_{1t})^2 = \frac{\mu_2}{N} + \sigma_Z^2 \quad (6)$$

$$E(m_{3t})^2 = \frac{\mu_6 + 9\mu_2^3 - 6\mu_4\mu_2}{N} \quad (7)$$

According to equation (5) the covariance between the first and third sample moment goes to zero when N increases to infinity. However, this does not guarantee that the correlation converges to zero as the variance of the third moment goes to zero, too. By substituting (5), (6) and (7) in (4) we see that the correlation coefficient will only tend to zero as N increases if $\sigma_Z^2 > 0$:

$$\rho_{1,3} = \frac{\left[1 - \frac{3}{N} + \frac{2}{N^2} \right] \mu_4}{\left[(\mu_2 + N\sigma_Z^2)(\mu_6 + 9\mu_2^3 - 6\mu_4\mu_2) \right]^{1/2}} \quad (8)$$

⁶ Bryan and Cecchetti (1999) obtain the excess fourth moment by subtracting 3 times the second moment (the kurtosis of the normal distribution) as last term in equation (5). This seems to be an error: direct calculation under the full independence assumption yields

$$E(m_{1t}m_{3t}) = E \left[\frac{1}{N} \sum ((x_{it} - Z_t) - Z_t) \left(\frac{1}{N} \sum ((x_{it} - Z_t) - \frac{1}{N} \sum (x_{it} - Z_t))^3 \right) \right] = \left[\frac{1}{N} - \frac{3}{N^2} + \frac{2}{N^3} \right] \mu_4$$

However, this difference is only of minor importance with our data as the fourth moment is very large.

Consequently, in small samples there is a bias, which can be substantial if the kurtosis is high. Intuitively, the problem arises because the population moments cannot be measured directly and have to be estimated. Thus, an extremely high or low observation on x_{it} leads to an artificial co-movement of the first and the third sample moment in small samples. This effect, which is of course stronger the higher the kurtosis is, is averaged out with increasing N as the sample mean and third moment converge to the population moments Z_t and μ_{3r} , respectively.⁷ This problem of course also affects a regression analysis of the relationship between the first and third moments. The slope of a simple regression of the first on the third moment is:

$$\beta_{1,3} = \frac{E(m_{1t}m_{3t})}{E(m_{3t})^2} = \frac{\left[1 - \frac{3}{N} + \frac{2}{N^2}\right]\mu_4}{\left[\mu_6 + 9\mu_2^3 - 6\mu_4\mu_2\right]} . \quad (9)$$

This regression coefficient does not converge to zero with increasing N under the assumption of independence over time and across observations.

Before turning to the consequences of this problem, we consider the relationship between the first and second moment if the cross-sectional distribution is asymmetric. Bryan and Cecchetti (1999) show that the correlation of mean and variance is

$$\rho_{1,2} = \frac{\left[1 - \frac{3}{N} + \frac{2}{N^2}\right]\mu_3}{\left[(\mu_2 + N\sigma_Z^2)(\mu_4 - \mu_2^2 - 6\mu_4\mu_2)\right]^{1/2}} \quad (11)$$

Thus, when the data are positively or negatively skewed, there is a bias that vanishes with N asymptotically if there is time variation in Z . The intuition is similar to that outlined for the mean and the third moment. If the data are positively (negatively) skewed, positive (negative) outliers lead to a positive (negative) spurious correlation between sample mean and variance which disappears with increasing N .⁸ This problem seems to be less serious than the case of the mean and the skewness for two reasons: First, the mean-skewness bias necessarily exists as the fourth moment is always nonzero whereas the problem for the mean-variance case does

⁷ Note that as N increases, the sample third moment converges to a constant (0 in our case of a symmetric distribution), whereas the sample mean converges to a time-varying variable since the variance of Z is positive. Thus in this case there is no correlation between the sample moments. If the variance of Z is zero, both series tend to constant values as N increases, and there is no identifiable relationship between them.

⁸ The consequences for regressions of the first and second moments are analogous to the first and third moments case discussed above and are not explicitly discussed here.

not arise with symmetric distributions. Secondly, the price-change distribution is sometimes (as in the US) nearly symmetrical but may have a very high kurtosis.

Bryan and Cecchetti demonstrate that this bias is empirically relevant in post-war US CPI and PPI data. They estimate the bias in the correlation between the mean and the skewness to be around 0.25 according to equation (8). This estimate, which is obtained under the independence assumption, should be considered as a lower bound since their Monte Carlo experiments showed that it is nearly twice as large when the auto and crosscorrelation properties of the components of the CPI are accounted for. Since the bias is potentially sizable, the problem needs to be taken into account in the econometric analysis below. We do this in two steps.

First we follow Bryan and Cecchetti and calculate the bias under the independence assumption by inserting the time-series means of the cross-section moments in equation (8) and (11). However, we use a procedure different from that of Bryan and Cecchetti (1999) to estimate the variance of Z . Instead of calculating the variance of high-order moving average, the long-run variance of the inflation rate m_{it} is obtained as estimate of the spectral density at frequency zero using a Bartlett kernel with bandwidth of 18.

It should be noted that with unequal weights on the individual subcomponents, the effective cross-section sample size is not simply equal to the number of subcomponents, N , since the calculation of the effective sample size must account for the differing weights. To see this, suppose that the CPI consists of two groups of goods: housing expenses, with a weight of 20 percent, and other goods with a weight of 80 percent. Even if the latter group can be disaggregated further, there will always be uncertainty associated with the former group. Thus, every component of the CPI with finite weight that cannot be disaggregated further leads to an effective finite sample size even if the number of other subcategories goes to infinity. The effective sample size is the inverse of the sum of the “statistical importance” of all observations (w_i^2). For equal weights $1/N$ we get, of course, an effective and nominal cross-section sample size equal to N . The effective sample size is obtained as $N = 1/\sum w_i^2$ and this value is used in the formulas (8) and (11). The resulting values, 10 for Hong Kong and 19 for Japan, are clearly lower than the number of components, 30 and 42, in the two data sets.

In Japan the correlation between the inflation rate and the standard deviation of the distribution of relative price changes is 0.258 in the inflation period (1982/1 – 1998/12) and

0.116 in the deflation period (1999/1 – 2004/6) and is thus relatively weak. The bias of the correlation between the mean and the variance seems to be non-negligible: using equation (11) and the time-series means of the cross-section moments and the long-run variance of inflation we get an estimate of the bias of 0.112 and 0.117 for the inflation and deflation periods.

With respect to the third moment, we note a high and positive correlation of 0.584 for the inflation, and 0.609 for the deflation, period. The correlation for the inflation period seems to be strongly distorted by the measurement problem discussed in Section 3, as the application of formula (8) yields a value for the correlation of 0.343 under the assumption of cross-section and time-series independence of inflation and relative price changes. Interestingly this bias problem is less important in the deflation period as the corresponding expected value of the correlation under the independence assumption is 0.112. The analogue correlations for Hong Kong for the periods 1982/1 – 1998/9 and 1998/10– 2003/6 are relatively low in all cases, except in that inflation and unexpected inflation is positively correlated with the third moment with a coefficient of 0.340 for the inflation and 0.342 for the deflation period. These values have, however, to be interpreted with care: the expected value obtained under the independence assumption is 0.264 (inflation) and 0.3263 (deflation). In sum, this correlation analysis indicates a potential bias problem in the regression estimated by Ball and Mankiw (1995).

The analogue correlations for Hong Kong for the periods 1981/7 – 1998/6 and 1998/7 – 2004/10 are relatively low in all cases, except in that inflation is positively correlated with the third moment with a coefficient of 0.340 for the inflation and 0.254 for the deflation period. These values have, however, to be interpreted with care: the expected value obtained under the independence assumption is 0.264 (inflation) and 0.305 (deflation). In sum, this correlation analysis indicates a potential bias problem in the regression estimated by Ball and Mankiw (1994 and 1995).

We therefore next seek to estimate the same regression in ways that avoid this bias. In an earlier version of the paper we used robust measures for the sample moments since these are less sensitive to outliers, which are the main source of the bias problem identified by Bryan and Cecchetti. However Monte Carlo experiments indicated that this approach did not fully mitigate the problem. In this version of the paper we use a random subsampling approach which proceeds as follows. For each month we generate for each subcategory of the CPI a random dummy variable which takes the value zero and one with equal probability. If this

variable is 1, we use the corresponding subcategory of the CPI to estimate the (weighted) mean of the distribution. If it is zero, we use it to calculate the standard deviation and skewness. This approach has the advantage that it creates no artificial correlation between the mean and the higher order moments since they are computed using randomly selected and different subsamples. We then use this measure to regress the mean on the standard deviation and mean. Of course the estimates obtained depend on the random sample selection. Thus, we run 1000 replications of this procedure and average the regression coefficients and their standard errors. In an appendix we present Monte Carlo evidence that this approach avoids the bias problem that arises in the case of OLS estimates.

4. Skewness, price dispersion and inflation in Japan and Hong Kong

Figures 2 and 3 show the first four (weighted) moments of the monthly rate of change (in percent) in the 42 and 30 seasonally adjusted sub-indices of the CPI of Japan and Hong Kong, respectively.⁹ Thus, besides the mean (CPI inflation), we plot the standard deviation (STD) and the coefficients of skewness (SKEW) and kurtosis (KURT), both of which are standardised. The sample periods span 1982/1 – 2006/4 for Japan and 1981/7 – 2004/10 for Hong Kong and thus exclude deliberately the high inflation environment of the seventies and the first years of the 1980s since we want to consider a period of moderate inflation and deflation.

--- Figures 2 and 3 here ---

The series displayed in Figures 2 and 3 are all rather volatile. Figure 2 does not suggest that there is a clear break in the mean of the price changes in the CPI subcategories in Japan around 1998/99 when the CPI started to decline. This suggests that tests of the menu cost model using data from the inflation and deflation periods in Japan may not be very different. By contrast, Figure 3 shows a break in the inflation rate of Hong Kong. Moreover, the volatility of the coefficients of skewness and kurtosis in Hong Kong seems to be lower in the deflation than the inflation period.

To obtain a more formal impression of the behaviour of inflation, next we follow Ball and Mankiw (1995) and regress inflation, measured by the mean of the distribution of individual

⁹ The reason we focus on sub-indices rather than at the individual components in the CPI is that the latter are difficult to model. For instance, in many cases prices are changed only rarely and then by large amounts. The discussion above regarding the effective cross-section sample size suggests that using subindices does not entail a large loss of information.

price changes, on the standard deviation and the skewness of that distribution. That is, we regress:

$$MEAN = \beta_1(1 - D_t) + \beta_2(1 - D_t)STD_t + \beta_3(1 - D_t)SKEW_t + \beta_4D_t + \beta_5D_tSTD_t + \beta_6D_tSKEW_t + \beta_7(1 - D_t)SKEW_t * STD_t + \beta_8D_tSKEW_t * STD_t + v_t$$

where D_t is a deflation dummy that takes the value of unity between 1999/1 – 2004/5 in Japan and between 1998/10 – 2003/6 in Hong Kong and zero otherwise.

Since preliminary regressions indicated that the residuals were heteroscedastic and that a large number of lags of the dependent variable were required to purge them of serial correlation, we decided to disregard the serial correlation in estimation but conduct inference using Newey-West standard errors, which are robust to serial correlation and heteroscedasticity.

Table 1 displays the OLS estimates using the full sample mean and higher-order moments. Although we know that these estimates can be strongly biased we report them for reasons of comparison.

--- Table 1 here ---

The results are discouraging in that all parameters are insignificant, except those on SKEW interacted with STD in Japan. Next we test the economically interesting hypothesis that the impact of STD on inflation is equal, but has opposite signs, in the inflation and deflation periods; and that the impact of SKEW and SKEW*STD is the same in both periods ($\beta_2 = -\beta_5$, $\beta_3 = \beta_6$, $\beta_7 = \beta_8$). Since the parameters are imprecisely estimated, it is not surprising that the test does not reject ($p = 0.206$ in Japan and $p = 0.808$ in Hong Kong).

Next we turn to the results of the approach using differing random subsamples for the calculation of the mean and the higher-order moments, respectively. The empirical results in Table 2 show the mean coefficients estimates obtained using 1000 replications.

--- Table 2 here ---

The results now indicate that SKEW*STD is significant and positive in Japan in both subperiods, as is SKEW in Hong Kong. Thus, the findings provide strong evidence in support of the menu cost model for both subperiods. However, the parameter estimates on the standard deviation of the distribution is insignificant in all cases, but the point estimate changes sign in Hong Kong. While this result is discouraging, the sample period is short and the change in the behaviour of inflation is similar in the two subsamples in Japan.

Since tests of the joint hypothesis that $\beta_2 = -\beta_5$, $\beta_3 = \beta_6$, $\beta_7 = \beta_8$ do not reject, next we impose these restrictions. Table 3 shows the resulting estimates.

-- Table 3 here ---

Not surprisingly the coefficients on SKEW in Hong Kong and SKEW*STD in Japan remain highly statistically significant. Moreover, the coefficient estimates for STD remain statistically insignificantly different from zero in both economies. Thus, the information gain coming from imposing the restrictions is not sufficient to render STD significant.

5. Conclusion

In this paper we explore the empirical validity of the menu cost model of Ball and Mankiw (1994 and 1995) and Lucas' (1973) island model. To do so, we study disaggregated monthly CPI data covering the last two decades for Japan and Hong Kong, which experienced inflation until the middle of 1998 and then underwent protracted deflation that ended between 2003 and 2004. Using data from periods with deflation is interesting since it allows us to explore whether the impact of price dispersion on inflation changes sign between inflation and deflation periods. This hypothesis, which is implied by the menu cost model, has apparently not been studied in the literature.

In designing the empirical framework we are mindful of the findings of Cecchetti and Bryan (1999) that show that parameters in regressions of the inflation rate on higher-order sample moments are in general biased in finite cross-sectional samples. In particular, high kurtosis of the relative-price change distribution may lead to a spurious relationship between inflation and skewness, either on its own (SKEW) or interacted with price dispersion (SKEW*STD). To mitigate this problem, for each month, we randomly assign the observations on the rate of price increase for the subcomponents of the CPI to one of two groups, and calculate the mean from one of these and the higher order moments of the distribution of price changes from the other.

The random sampling procedure yields positive and highly significant coefficient estimates on SKEW or SKEW*STD in both periods and both countries. This finding is supportive of the menu cost model but difficult to rationalise using the Lucas model. However, the coefficient the standard deviation (STD) is insignificant for both economies and in both periods. While this finding runs counter to the predictions of the menu cost model in the presence of trend inflation or deflation, it is compatible with the model in an environment when the economy is close to price stability. To test this particular implication of the menu cost model, it would be

of interest to use data in which deflation was longer lasting and more rapid. That we leave for another paper.

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Figure 1
Consumer Prices (Changes over 4 quarters)

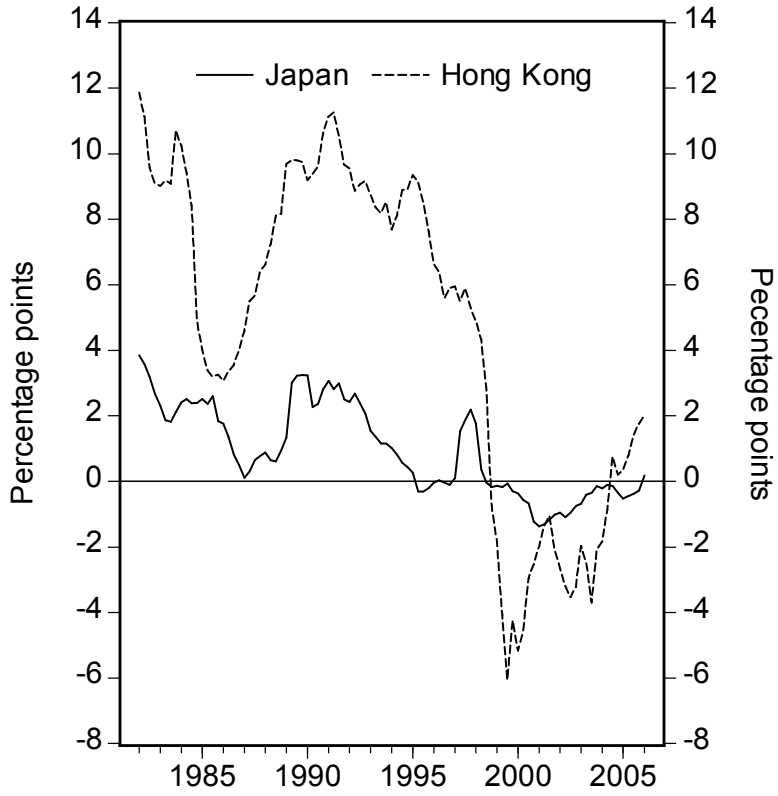


Figure 2: Moments of monthly changes in 42 CPI sub-indexes, Japan 1982/1-2006/4

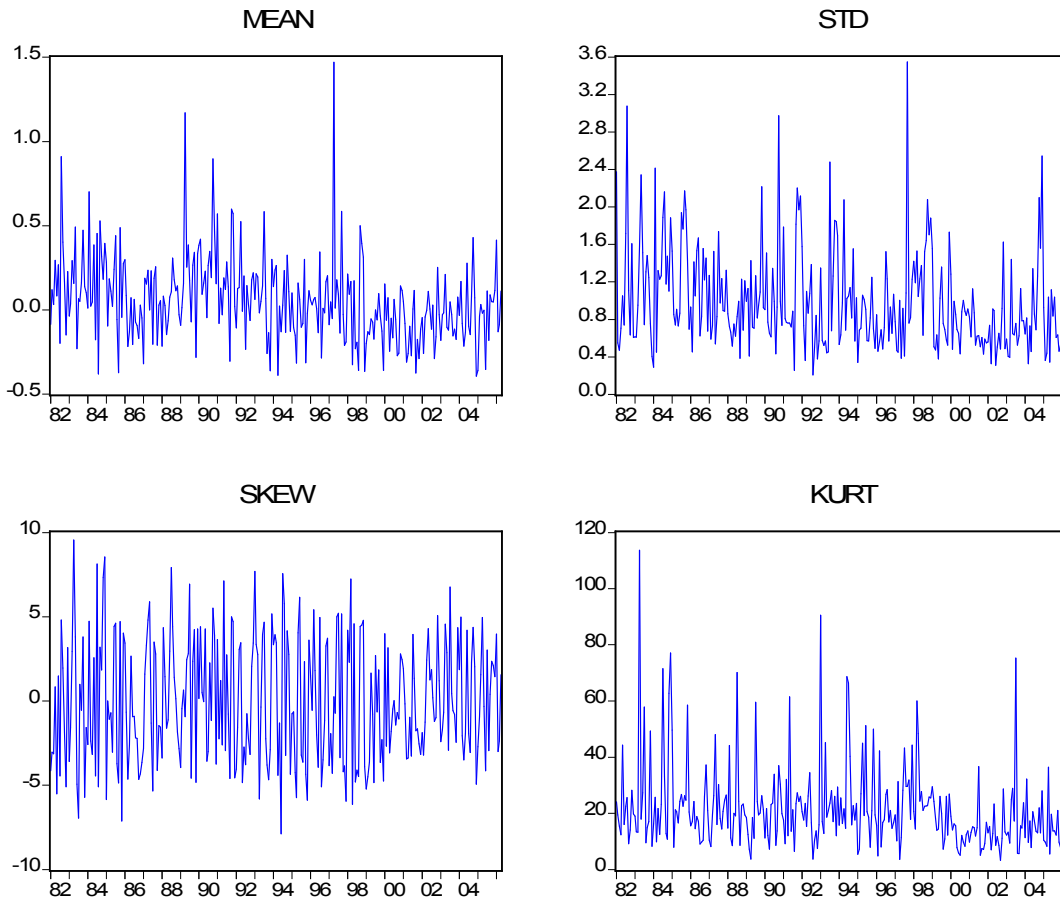


Figure 3: Moments of monthly changes in 30 CPI sub-indexes, Hong Kong 1981/7-2004/10

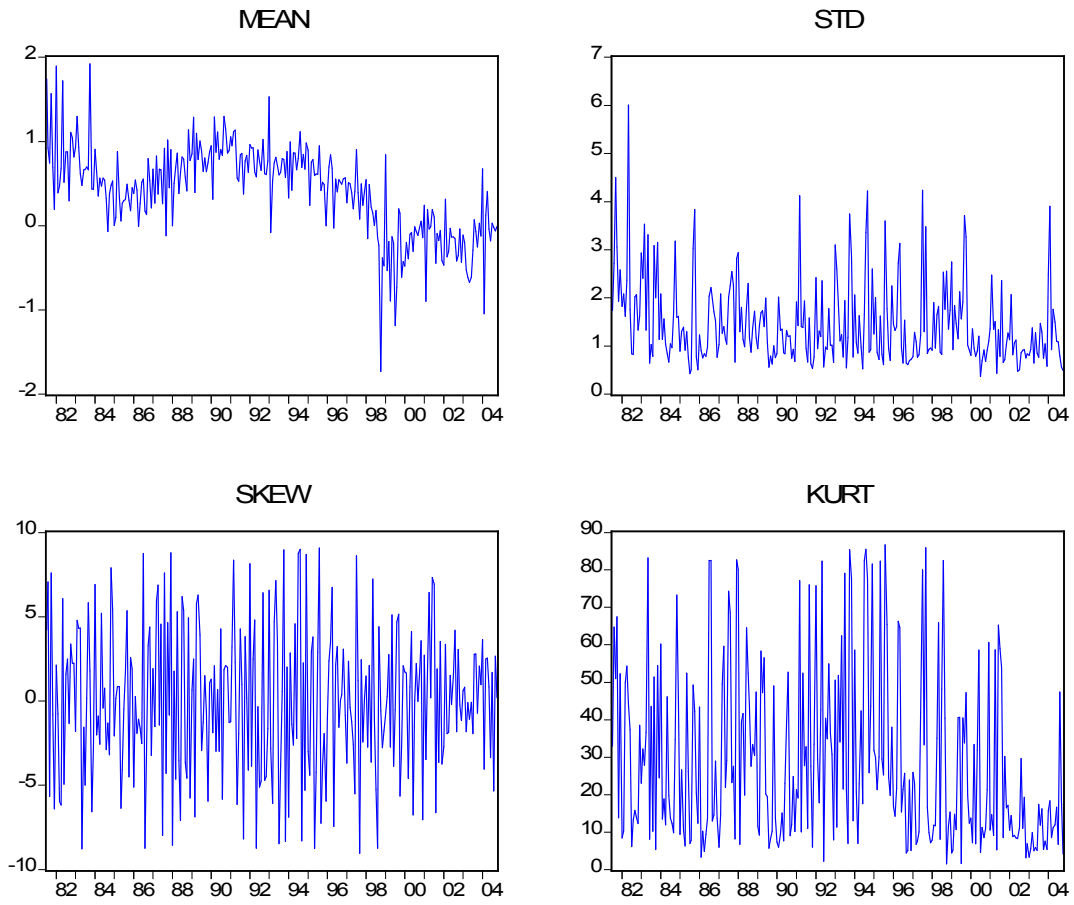


Table 1:
OLS regression estimates with Newey-West standard errors
 Full samples for mean and std/skew
 Standard errors of regression coefficient estimates in parentheses

$$\text{Mean} = \beta_1(1 - D_t) + \beta_2(1 - D_t)\text{STD}_t + \beta_3(1 - D_t)\text{SKEW}_t + \beta_4 D_t + \beta_5 D_t \text{STD}_t + \beta_6 D_t \text{SKEW}_t + \beta_7(1 - D_t)\text{SKEW}_t * \text{STD}_t + \beta_8 D_t \text{SKEW}_t * \text{STD}_t + v_t$$

D_t : Deflation dummy

	Japan	Hong Kong
STD (Inflation)	0.021	-0.007
	(0.019)	(0.027)
STD (Deflation)	-0.003	0.005
	(0.025)	(0.058)
SKEW (Inflation)	-0.010	0.018
	(0.006)	(0.012)
SKEW (Deflation)	0.013	0.038
	(0.010)	(0.025)
SKEW *STD (Inflation)	0.034***	0.007
	(0.005)	(0.005)
SKEW *STD (Deflation)	0.028***	-0.007
	(0.008)	(0.014)
Test of difference between inflation/deflation slopes (χ^2_3) $\beta_2 = -\beta_5, \beta_3 = \beta_6, \beta_7 = \beta_8$	4.567 $p\text{-value} = 0.206$	0.972 $p\text{-value} = 0.808$
Adjusted R-squared	0.559	0.635
Durbin-Watson stat.	1.854	1.527

Note: ***/** denotes significance at 10/5/1 percent level. For Japan the inflation sample is 1982/1 – 1998/12 and the deflation sample is 1999/1 - 2004/5; for Hong Kong the two samples are 1982/1 - 1998/9 and 1998/10 - 2003/6.

Table 2:
OLS sampling regression estimates
 Randomly selected subsamples for mean and std/skew
 Mean of 1000 Monte Carlo replications
 Standard errors of regression coefficient estimates in parentheses

$$Mean = \beta_1(1 - D_t) + \beta_2(1 - D_t)STD_t + \beta_3(1 - D_t)SKEW_t + \beta_4D_t + \beta_5D_tSTD_t + \beta_6D_tSKEW_t + \beta_7(1 - D_t)SKEW_t * STD_t + \beta_8D_tSKEW_t * STD_t + v_t$$

D_t : Deflation dummy

	Japan	Hong Kong
STD (Inflation)	0.026 (0.024)	0.030 (0.024)
STD (Deflation)	0.003 (0.052)	-0.018 (0.096)
SKEW (Inflation)	0.090 (0.130)	0.049*** (0.019)
SKEW (Deflation)	0.320 (0.237)	0.049*** (0.017)
SKEW *STD (Inflation)	0.033*** (0.009)	0.005 (0.007)
SKEW *STD (Deflation)	0.029* (0.016)	0.008 (0.015)
Test of difference between inflation/deflation slopes (χ^2_3) $\beta_2 = -\beta_5, \beta_3 = \beta_6, \beta_7 = \beta_8$	3.146 p -value: = 0.370	2.802 p -value = 0.423
Adjusted R-squared	0.398	0.447
Durbin-Watson stat.	1.879	1.706

Note: ***/*** denotes significance at 10/5/1 percent level. For Japan the inflation sample is 1982/1 – 1998/12 and the deflation sample is 1999/1 - 2004/5; for Hong Kong the two samples are 1982/1 - 1998/9 and 1998/10 - 2003/6.

Table 3:
OLS sampling regression estimates
 Randomly selected subsamples for mean and std/skew
 Mean of 1000 Monte Carlo replications
 Standard errors of regression coefficient estimates in parentheses

$$Mean = \beta_1(1 - D_t) + \beta_2(1 - D_t)STD_t + \beta_3(1 - D_t)SKEW_t + \beta_4D_t - \beta_2D_tSTD_t + \beta_3D_tSKEW_t + \beta_7(1 - D_t)SKEW_t * STD_t + \beta_8D_tSKEW_t * STD_t + v_t$$

D_t : Deflation dummy

	Japan	Hong Kong
STD	0.023	0.029
	(0.022)	(0.025)
SKEW	0.121	0.054***
	(0.129)	(0.016)
SKEW *STD	0.032***	0.004
	(0.009)	(0.007)
Adjusted R-squared	0.396	0.444
Durbin-Watson stat.	1.877	1.706

Note: ***/** denotes significance at 10/5/1 percent level. For Japan the inflation sample is 1982/1 – 1998/12 and the deflation sample is 1999/1 - 2004/5; for Hong Kong the two samples are 1982/1 - 1998/9 and 1998/10 - 2003/6.

Appendix: Some Monte Carlo Results

In this appendix we report the results of a small Monte Carlo study in which we explore whether the random sampling estimating strategy is sensitive to the bias identified by Bryan and Cecchetti (1999). For reasons of comparison we perform the analysis for the full sample mean and higher-order moments. We first use the mean and the usual higher order moments before replacing the mean by the robust median.

We consider a panel with cross-section dimension N and time dimension T . We assume that the data, denoted by x_{it} , are driven solely by one common element, namely a time-varying normally distributed mean Z (that is, $\sigma_Z^2 > 0$). Formally:

$$\begin{aligned} Z_t &= E_i x_{it} \\ E(Z_t) &= 0 \\ E(Z_t)^2 &= \sigma_Z^2, \\ E(Z_t Z_{t-s}) &= 0, \quad s \neq 0 \end{aligned}$$

The x_{it} are identically and independently distributed with mean 0 and variance 1 across i and stationary over t . Corresponding to the data used in this paper the variance of Z is assumed to be 0.25, N is equal to 40 and T is set to 200. 1000 Monte Carlo replications of such samples were generated for three different distributions of x_{it} : a symmetric fat tailed t -distribution and a asymmetric (standardized) χ^2 -distribution with 4 and 10 degrees of freedom, respectively. The type of the distributions was selected since the bias problem for SKEW is particularly important for fat tailed distribution whereas the bias for STD is mainly present with asymmetric distributions.

Then the relationship between the sample mean and the sample standard deviation and skewness is analyzed by running a regression of the mean on the two higher-order moments. In Table A1 we show the mean and standard deviations of the regression coefficients for STD and SKEW. The results obtained by using the mean and the higher-order moments for all three measures is documented in Table A2 and the result of the specification replacing the mean by the median are presented in Table A3.

Table A1 shows that the random sampling approach works: the means of the regression coefficients of the 1000 replications are very small not statistically different from zero. Thus this approach provides us with unbiased regression estimates since by construction there is no relationship between mean and higher order moments in these samples. By contrast, using the

original mean and higher-order moments leads to sometimes strongly biased estimates, as indicated in Table A2. The bias is particularly large for the asymmetric distributions, where the coefficients of SKEW and STD are both heavily biased. According to Table A3 these biases cannot be avoided by using the median as regressand. The same applies if we use robust measures for the second and third moment as regressors.

Table A1: Mean and standard deviation (in parentheses) of regression coefficients of random subsamples for mean and STD/SKEW, N=40, T=200, different i.i.d. cases, 1000 replications

Distribution	STD	SKEW
t , 4 df	0.000	0.000
	(0.004)	(0.001)
t , 10 df	0.002	-0.001
	(0.004)	(0.001)
χ^2 , 4 df	-0.001	0.000
	(0.004)	(0.002)
χ^2 , 10 df	0.006	-0.002
	(0.005)	(0.001)

Table A2: Mean and standard deviation (in parentheses) of regression coefficients of full sample mean on STD/SKEW, N=40, T=200, different i.i.d. cases, 1000 replications

Distribution	STD	SKEW
t , 4 df	-0.011	0.079
	(0.004)	(0.001)
t , 10 df	-0.002	0.041
	(0.005)	(0.002)
χ^2 , 4 df	0.782	-0.121
	(0.005)	(0.002)
χ^2 , 10 df	0.581	-0.072
	(0.006)	(0.002)

Table A3: Mean and standard deviation (in parentheses) of regression coefficients of full sample for median on STD/SKEW, N=40, T=200, different i.i.d. cases, 1000 replications

Distribution	STD	SKEW
t , 4 df	-0.008	-0.042
	(0.004)	(0.001)
t , 10 df	-0.004	-0.112
	(0.005)	(0.002)
χ^2 , 4 df	0.503	-0.189
	(0.005)	(0.003)
χ^2 , 10 df	0.423	-0.194
	(0.006)	(0.002)