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Evidence from Japan and Hong Kong

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September 1, 2007

Abstract

We test the menu cost model of Ball and Mankiw (1994, 1995), which implies that the impact of price dispersion on inflation should differ between inflation and deflation episodes, using data for Japan and Hong Kong. We use a random cross-section sample split when calculating the moments of the distribution of price changes to mitigate the small-cross-section-sample bias noted by Cecchetti and Bryan (1999). The parameter on the third moment is positive and significant in both countries during both the inflation and deflation periods, and the parameter on the second moment changes sign in the deflation period, as the theory predicts.

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

JEL Numbers: E31

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[*] We are grateful to seminar participants at University of Basel, the Swiss National Bank, University of Bern, and Deutsche Bundesbank; and to Katrin Assenmacher-Wesche, Richhild Moessner and Ulf von Kalkreuth (who suggested the use the sample split method to calculate the mean and the higher moments of the distribution) for helpful comments.

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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 H. 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 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 exogenous 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 infla-

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[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.

tion drives relative prices apart is that 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 skew 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 remain 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 chose to do so, while the firms that would like to cut prices a little will refrain from doing so. As a consequence, positive skew will be associated with rising, and negative skew with declining, prices.

While this argument suggests that solely skew is important in the determination of inflation, the dispersion of prices matters 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 skew. 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 in-

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

crease 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 will raise the likelihood that average prices are falling.

Second, suppose that the trend rate of inflation is positive and that the distribution of price changes is symmetric. As shown by Ball and Mankiw (1994 and 1995), 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 declines will not change nominal prices but merely wait and let inflation reduce their relative price gradually to the appropriate level. Firms who have experienced a positive shock to their equilibrium prices can not rely on this adjustment mechanism and therefore 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.

It is important to note that while skew 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 their implications for skew. First, the Ball-Mankiw model implies that the relationship between inflation and relative price variability relationship should disappear if non-standardized skewness (that is, skewness multiplied by the standard deviation of relative price changes) is included in the regression whereas the Lucas model has no such implication. Second, in the Ball-Mankiw model the correlation between inflation and skewness should be positive irrespectively of whether the economy is an inflationary or deflationary environment. Again, the Lucas model makes no such prediction. By contrast, if the distribution of relative price shocks are symmetric (so that skew is zero) and the trend rate of inflation is positive, however, it is impossible to distinguish between the two models since they both imply that an increase of the standard deviation of the distribution of individual price changes raises inflation .

This paper expands the existing literature in two ways. First and as noted above, it is of interest to test the Ball-Mankiw model using data from episodes trend deflation. 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. These data thus allow us to test the menu cost model by exploring whether the sign of the impact of the dispersion of relative prices depends on whether the economy is experiencing trend inflation or trend deflation.

Second, we propose a method to overcome the problem identified by Stephen G. Cecchetti and F. Michael 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 inflation and 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 the small-cross-section-sample bias. Furthermore, we shat that the bias problem is mitigated if the variance of the long-run or trend inflation rate is high, which is an additional reasons for why the use of inflation and deflation period data (which contain considerable long-run variation of the inflation rate which should mitigate the bias problem) 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 skew 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 skew in both periods and both countries. Furthermore, the coefficient the standard deviation is positive and highly significant in the inflation period but negative, but less significant, in the deflation period. These results provide strong support for the menu cost hypothesis. 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 several 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, peaked in the early 1990s. It subsequently declined before turning to deflation in around the middle of 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. 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 depreciation 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. Despite the difference in the severity of the episode, it is clear that deflation was deeply entrenched in both economies.

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

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[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.

The data analysed below consists 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:

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

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:

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

The standard deviation (STD) as well as the coefficients of skewness (SKEW) and kurtosis (KURT) are the 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 strongly affected 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. Below we 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 $x_{it} \equiv \Delta p_{it}$), are driven solely by one common element, namely a time-varying mean *Z* (that is, $\sigma_Z^2 > 0$). Formally:

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

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:

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

(4)

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

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

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

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

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.

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[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

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

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

By substituting (5), (6) and (7) in (4) we see that the correlation coefficient will only tend to zero as N increases if $\sigma_Z^2 > 0$:

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

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:

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

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 large T correlation of mean and variance is

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

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 as 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

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[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.

between sample mean and variance which disappears under the assumption of time variation of the population mean with increasing *N*. This problem seems to be less serious than the case of the mean and the skew for two reasons: First, the mean-skew bias necessarily exists as the fourth moment is always nonzero whereas the problem for the mean-variance case does not arise with symmetric distributions. Secondly, the price-change distribution is sometimes (as in the US) nearly symmetrical but have a very high kurtosis. Of course, 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.

Bryan and Cecchetti demonstrate that this bias is empirically relevant in post-war US CPI and PPI data. They estimated the bias in the correlation between the mean and the skew to be around 0.25 according to equation (8). This estimate, which obtained under the independence assumption, should be considered as a lower bound as their Monte Carlo experiments showed that it is nearly double as large when the auto and cross correlation 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 large-*T* 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_{1t} 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 understand 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 and 19, 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 0.123 in the deflation period (1999/1 – 2006/4) and are 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 0.197 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 negligible in the deflation period as the corresponding expected value of the correlation under the independence assumption is 0.073. 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 and unexpected 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 (1995).

We therefore turn to the second step of our approach in which we 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 some Monte Carlo experiments reported partly in an appendix indicated

that this approach did not mitigate the problem. In this version of the paper we use a random subsampling approach which proceeds as follows. For each month we generate a random dummy variable which takes the value zero and one with equal probability with sample size N. 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 skew. 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. Skew, 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.[8] 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. 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.

--- 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 break in the series around 1997 when the Japanese CPI started to decline. By contrast, Figure 2 clearly 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.

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[8] 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.

To obtain a more formal impression of the behaviour of inflation, next we report the results of OLS regressions of inflation on the standard deviation and the skew of the distribution of price changes. 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. We estimate this equation for inflation and deflation periods using a dummy variable that take the value zero during the inflation period and one in the deflation period. For Japan the two periods are defined as1982/1 – 1998/12 and 1999/1 – 2006/4; for Hong Kong they are 1981/8 – 1998/6 and 1998/7 – 2004/10. 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 coefficient of SKEW is positive and highly statistically significant for both countries and both periods. Moreover, there is no statistically significant difference between the estimates for the inflation and deflation periods. The coefficient of STD is statistically not different from zero except for Japan in the inflation period.

Now let us 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 using 1000 replications.

--- Table 2 here ---

The parameter on the standard deviation of the distribution is statistically insignificantly different from zero in both countries and both subperiods, suggesting the importance of accounting for the bias problem given the significant correlations of full sample moments reported above. For Hong Kong we find the change of a positive to a negative coefficient from the inflation to the deflation period, as the Ball Mankiw model would let us to expect. We obtain a positive and highly significant coefficient on skew in both periods and both countries and the hypothesis that there is no change in the slope parameters between the inflation and deflation period cannot be rejected. Thus, we conclude that the empirical analysis supports partly, i.e. with respect to SKEW the menu cost explanation for the relation between inflation and the distribution of relative prices changes. However the problem is that the standard de-

viation coefficient is statistically insignificant. In order to get more precise estimates we adopted some restrictions across periods and countries. First, we assume that the coefficients for the standard deviation is equal across countries in both periods. Second, we adopt the restriction that the skewness coefficient is in both countries the same across periods. Table 2 shows the restricted SUR estimates of this model.

-- Table 3 here ---

Unfortunately the coefficient estimates for the STD coefficients remain statistically insignificantly different from zero although the sign on the constrained coefficient on the standard deviation changes sign between the inflation and the deflation periods. Not surprisingly the coefficient on SKEW remains highly statistically significant in both countries. Moreover, the parameter restrictions are not rejected by the data.

5. Conclusion

In this paper 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 2004 and 2005. These data thus allow us to explore whether and, if so, how the relationship between changes in absolute price level and changes in the distribution, in particular the standard deviation, of relative prices varies between episodes of trend inflation and trend deflation, an issue which has apparently not been studied in the literature.

In designing the empirical framework we are mindful of the findings of Cecchetti and Bryan (1999) who 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. 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 empirical results of the random sampling procedure yields a positive and highly significant coefficient on skew in both periods and both countries. However, the coefficient the standard deviation is insignificant both subperiods and we note the ex-

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pected change in sign only for Hong Kong data. Adopting cross country (common coefficient for the standard devaiation across countries) and cross period restrictions (common skew coefficient across period in both countries) in a SUR framework leads to a change in sign of the standard deviation coefficient, although the result lacks statistical significance. Overall we conclude that the empirical analysis provides result which are in line with the menu cost explanation for the relation between inflation and the distribution of relative prices changes.

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Figure 1
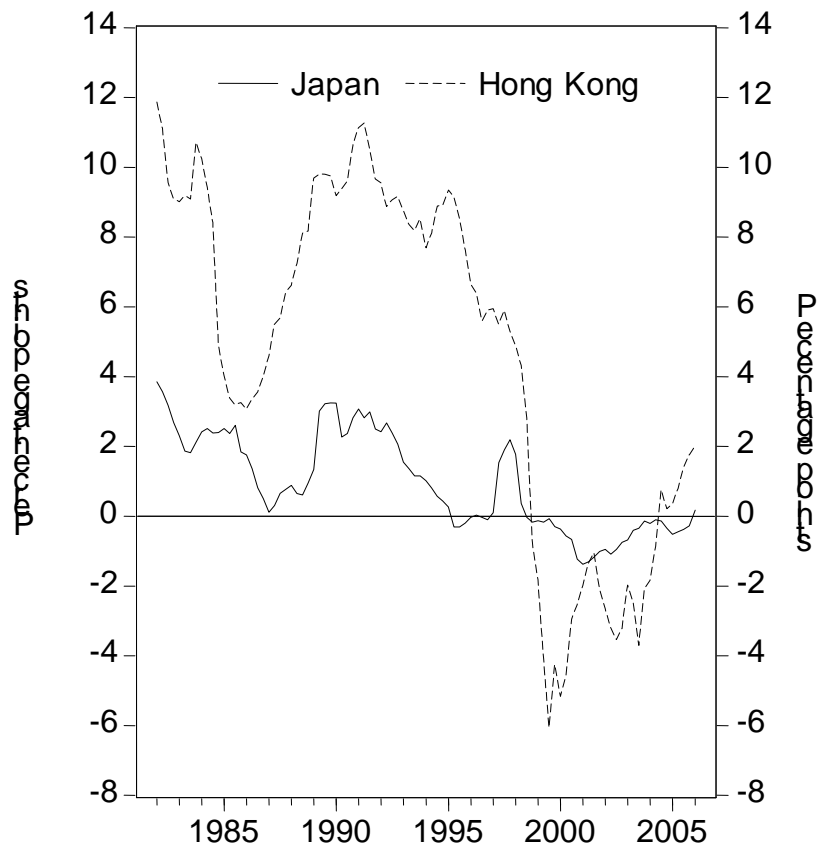Consumer Prices (Changes over 4 quarters)

Figure 2: Moments of monthly changes in 42 CPI sub indexes, Japan 1982/1-2006/4
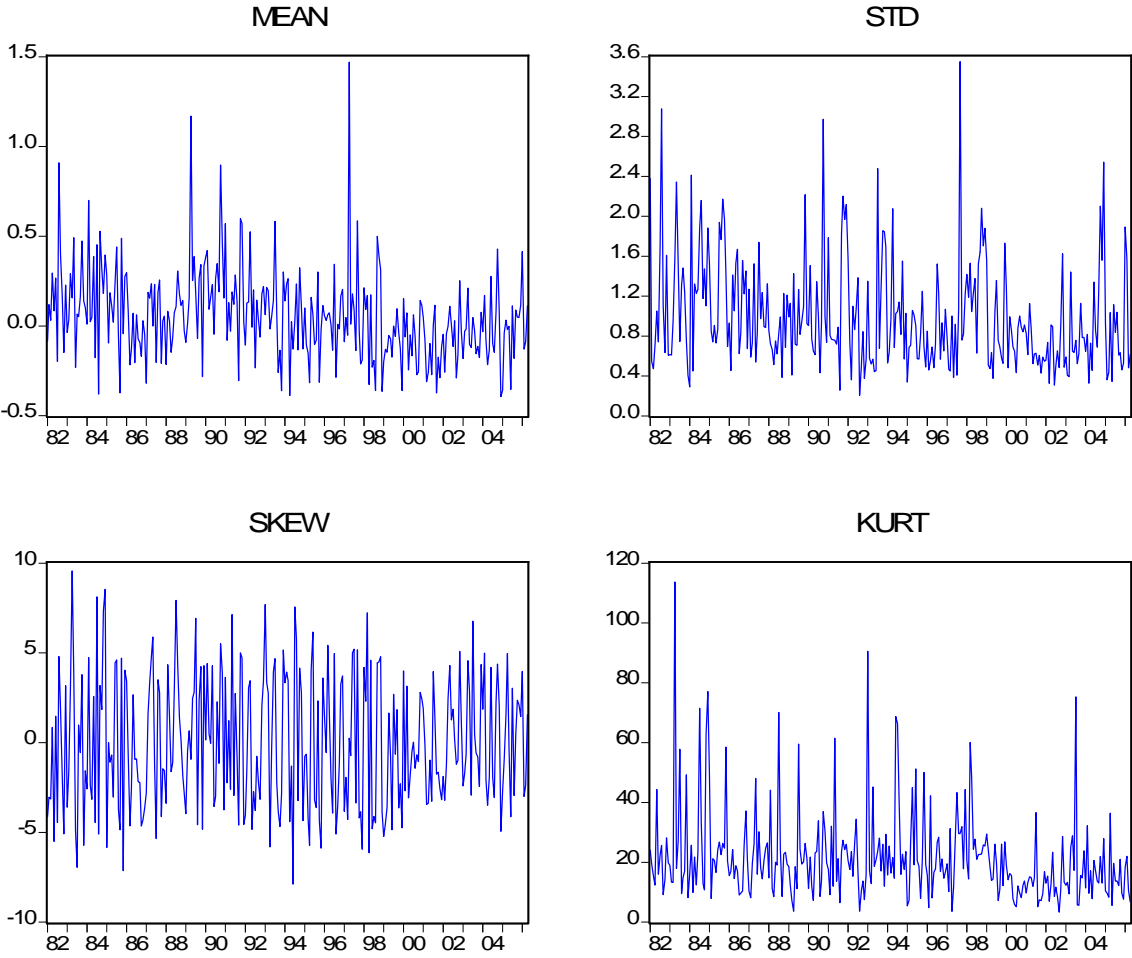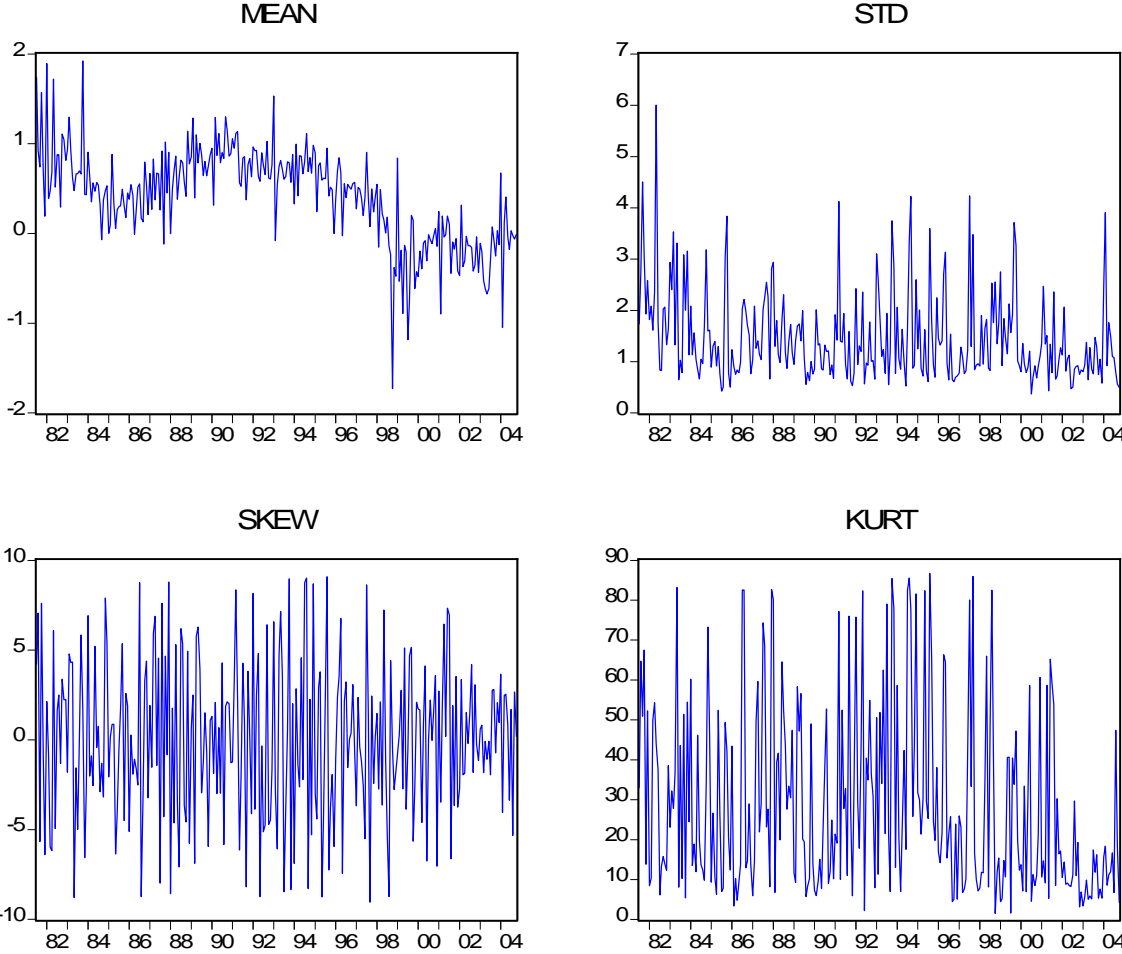
Figure 3: Moments of monthly changes in 30 CPI sub indexes, Hong Kong 1981/7-2004/10

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

$$Mean = \beta_1(1 - D_t) + \beta_2(1 - D_t)STD_t + \beta_3(1 - D_t)SKEW_t + \beta_4 D_t + \beta_5 D_t STD_t + \beta_6 D_t SKEW_t$$

D_t : *Deflation dummy*

	Japan	**Hong Kong**
STD (Inflation)	0.0609**	0.0387
	(0.0266)	(0.0263)
STD (Deflation)	0.0104	-0.0198
	(0.0309)	(0.0709)
SKEW (Inflation)	0.0409***	0.0362***
	(0.0036)	(0.00040)
SKEW (Deflation)	0.0460***	0.0346***
	(0.0049)	(0.0129)
Test of difference between inflation/deflation slopes (χ_2^2)	1.3965	2.3979
Adjusted R-squared	0.4988	0.6239
Durbin-Watson stat	1.8608	1.6070

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 - 2006/4; for Hong Kong the two samples are 1982/1 - 1998/6 and 1998/7 - 2004/10.

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

$$Mean = \beta_1(1-D_t) + \beta_2(1-D_t)STD_t + \beta_3(1-D_t)SKEW_t + \beta_4 D_t + \beta_5 D_t STD_t + \beta_6 D_t SKEW_t$$

D_t : Deflation dummy

	Japan	**Hong Kong**
STD (Inflation)	0.0448	0.0378
	(0.0377)	(0.0280)
STD (Deflation)	0.0094	-0.0331
	(0.0417)	(0.0744)
SKEW (Inflation)	0.6513***	0.0648***
	(0.0781)	(0.0102)
SKEW (Deflation)	0.7530***	0.0075***
	(0.1238)	(0.0245)
Test of difference between inflation/deflation slopes (χ_2^2)	2.809	4.4802
Adjusted R-squared	0.3708	0.4662
Durbin-Watson stat	1.9226	1.8834

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 - 2006/4; for Hong Kong the two samples are 1982/1 - 1998/6 and 1998/7 - 2004/10.

Table 3: SUR estimates with restrictions
SKEW coefficients equal across inflation/deflation peridods
STD coefficients equal across countries

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

	Japan	**Hong Kong**
STD (Inflation)	0.0321	0.03205
	(0.0193)	(0.0193)
STD (Deflation)	-0.0089	-0.0089
	(0.0448)	(0.0448)
SKEW (Inflation)	0.6761***	0.0151***
	(0.0579)	(0.0041)
SKEW (Deflation)	0.6761***	0.0151***
	(0.0579)	(0.0041)
Test of restrictions (χ_4^2)	7.6124	
Adjusted R-squared	0.3203	0.3582
Durbin-Watson stat	1.9100	1.8760

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 - 2004/10; for Hong Kong the two samples are 1982/1 - 1998/6 and 1998/7 - 2004/10.

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:

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

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 four 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. Then the relationship between the sample mean and the sample standard deviation and skew 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 full cross section sample moments for 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. The type of the distributions were 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.

Table A1 shows that the subsample 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 as by construc-

tion there is no relationship between mean and higher order moments in these samples. By contrast using full cross section sample for the mean and the 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 coefficient 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	0.0003288 (0.004014)	0.0003383 (0.001294)
t, 10 df	0.002491 (0.003572)	-0.001002 (0.001262)
χ^2 ,4 df	-0.0009363 (0.004176)	0.0003225 (0.001508)
χ^2 , 10 df	0.005999 (0.004547)	-0.001603 (0.001483)

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	-0.01056 (0.004092)	0.07912 (0.001115)
t, 10 df	-0.001525 (0.005220)	0.04052 (0.001528)
χ^2 ,4 df	0.78192 (0.004945)	-0.1205 (0.002342)
χ^2 , 10 df	0.58050 (0.005738)	-0.07168 (0.001768)

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	-0.008154	(0.004277)	-0.04174	(0.001205)
t, 10 df	-0.003961	(0.005443)	-0.11161	(0.001593)
χ^2, 4 df	0.50254	(0.005432)	-0.18939	(0.002739)
χ^2, 10 df	0.42295	(0.006177)	-0.19396	(0.001892)

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