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Reconsidering the Relationship between Inflation and Relative Price Variability

It has long been popularly believed that the relationship between inflation and relative price variability (RPV) is positive and stable. Using disaggregated CPI data for the United States and Japan, however, this study finds that the relationship is neither linear nor stable over time. The overall relationship is approximately U-shaped around a nonzero threshold inflation rate. RPV therefore changes not with the inflation rate *per se*, but with the deviation of inflation from the threshold inflation rate. More importantly, the relationship is by no means stable over time but instead varies significantly in a way that coincides with regime changes of inflation or monetary policy. The relationship was positive during the period of high inflation of the 1970s and the early 1980s, as has been documented by a number of previous studies, whereas it takes a U-shape profile during the Great Moderation. The results are robust to the use of core inflation, which excludes the traditionally volatile prices of food and energy. This paper then presents a modified version of the Calvo-type sticky price model to describe the observed empirical regularities. Simulation experiments show that the modified Calvo model fits the data well, and that the underlying relationship hinges upon the degree of price rigidity, which is systematically related to inflation regime. For countries and periods with low inflation rates, the relationship takes a U-shape as price adjustment is more sticky. In a high-inflation environment, when price setting becomes more flexible, the U-shaped profile vanishes.

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THIS PAPER INVESTIGATES the relationship between inflation and relative price variability (RPV).[1] Because of its important implications for the welfare costs of inflation and the theory of monetary neutrality, this relationship has been a central theme of modern macroeconomics, receiving substantial attention in both theoretical and empirical research. The theoretical models have generally predicted a positive association between these two variables and this linkage has been supported by a large body of empirical evidence from various countries for different time periods.[2]

Despite the large literature, however, the connection is not fully understood on a couple of grounds. First, there is little consensus on the underlying functional form of the relationship. The empirical work that provides the prima facie evidence of a positive relationship between RPV and inflation has generally utilized linear regression models. The linearity assumption, however, has often been called into question by studies that provide evidence that the relationship is quadratic or piecewise linear (e.g., Parks 1978, Hartman 1991, Reinsdorf 1994). This includes a recent study by Fielding and Mizen (2008) who find a U-shaped relationship using quarterly U.S. Personal Consumer Expenditure (PCE) data for the period of 1967–2003. Identifying the correct functional form of the relationship has crucial implications for monetary policy. If the true relationship is positive, monetary authorities can reduce RPV by lowering inflation via disinflationary policy. This is no longer the case if the relationship is nonmonotonic.

Second, very little is known about the stability of the relationship. This is primarily because the existing literature has generally utilized linear regression analyses that assume the marginal impact of inflation on RPV is time invariant. The empirical evidence on structural changes of inflation series in numerous countries (e.g., Cogley and Sargent 2005, Levin and Piger 2004), however, provides ample reasons to doubt that the relationship remains unchanged across different regimes of inflation or monetary policy. Indeed, recent studies by Dabús (2000) and Caglayan and Filiztekin (2003) report that the relationship between inflation and RPV differs in an important manner, depending on the average inflation rate.[3]

The principal objective of this paper is to reexamine the link between inflation and RPV with an emphasis on understanding the functional form of the relationship

1. Throughout the paper, RPV is measured as the standard deviation (s.d.) of sectoral inflation rates relative to the aggregate rate. See Section 1.1 for a formal definition.

2. Popular economic models, such as the menu cost and the Lucas-type incomplete information approaches, typically predict a positive relationship in a positive inflation environment and this is generally supported by a considerable empirical literature beginning with the work of Vining and Elwertowski (1976). In this vein, the theoretical models are said to be "observationally equivalent" in terms of their predictions regarding the inflation-RPV linkage (Parsley 1996). See Lastrapes (2006) for more recent empirical evidence on positive relationship. For a few dissenting views, however, the reader is referred to the studies by Konieczny and Skrzypacz (2005), Reinsdorf (1994), and Silver and Ioannidis (2001) who report a negative relationship.

3. Dabús (2000) finds that the relationship between inflation and RPV in Argentina exhibits structural changes across different levels of inflation. Using price data from Turkey, Caglayan and Filiztekin (2003) also show that the association between inflation and RPV is significantly different between low and high inflationary periods.

and its stability over time. Instead of exploring the functional form and its stability separately, I analyze them jointly as the interface between them is believed to be essential for a proper understanding of the inflation-RPV nexus. To my knowledge, this is the first systematic attempt at such an analysis. To this end, I consider the case of two major economies, the United States and Japan, that have maintained relatively stable and low inflation rates during the Great Moderation period starting in the mid-1980s. Since my data set starts from the 1970s, these data span diverse regimes of inflation and monetary policy. This includes the deflationary episode of Japan in the mid-1990s that has rarely been studied in the literature.

This study offers interesting new theoretical and empirical insights on the relationship between inflation and RPV. On the empirical side, I find that the overall relationship takes a U-shape profile as in Fielding and Mizen (2008). More notably, the relationship is far from stable over time and instead varies significantly across sample periods. This is demonstrated in Figure 1, which displays scatterplots of inflation and RPV in the United States and Japan for the full sample period and for three subsample periods, using 1984 and 1997 as breakpoints.[4] As shown in the figure, in most cases, the nonlinear specification considered in this paper provides a superior characterization of the association between inflation and RPV relative to linear models. More specifically, the overall relationship is best described as U-shaped, as RPV dips with higher levels of inflation initially but begins to rise back when inflation increases gradually after passing a certain threshold. The U-shaped relationship implies that the marginal effect of inflation on RPV differs according to the level of inflation. For inflation levels below a threshold, higher inflation leads to lower RPV, while for inflation levels above the threshold, inflation increases RPV. From a policy perspective, this implies that disinflationary policies may not improve welfare if the benefit from lower inflation is outweighed by the costs from increased volatility of relative prices. A U-shaped relationship is observed not just in the United States, in accordance with Fielding and Mizen (2008), but also in Japan particularly during its deflationary period. Using core inflation, which strips out traditionally volatile prices such as those of food and energy items, does not alter these results.[5]

The empirical evidence in this paper also suggests that this U-shaped relationship is not stable over time. Instead, the relationship varies across sample periods. The U-shaped pattern observed for both countries during the Great Moderation when inflation stabilized substantially is not as apparent for the high-inflation episodes that

4. Year 1984 marks the onset of the so-called "Great Moderation" period when the volatility of aggregate economic variables, including inflation, declined significantly in most industrial countries, whereas 1997 is the year when the Japanese CPI began a downward trend. Though Japan experienced declining prices between the middle of 1995 and the end of 1996, the continuous price decline did not begin until 1997. The choice of break points seems admittedly somewhat arbitrary, but it is supported by more formal econometric analysis in Section 2.

5. A number of earlier studies that found a positive relationship between inflation and RPV were conducted on data from high-inflation periods, such as the oil shock periods in the mid-1970s and early 1980s. Some authors (e.g., Fischer 1981, Hartman 1991) attribute the positive relationship to an artifact of the higher inflation episodes of this period that may have resulted from common supply shocks, such as food and energy prices.

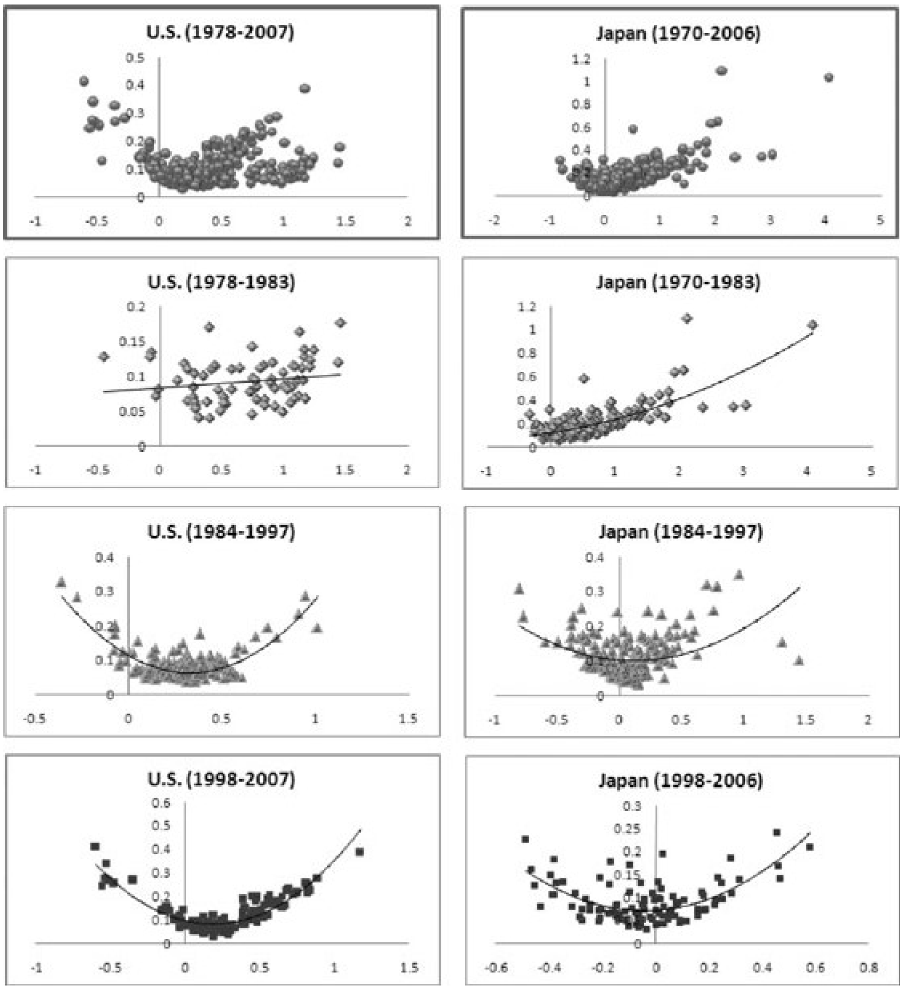

FIG. 1. Scatterplots of Relative Price Variability (Vertical Axes) and Monthly Inflation (Horizontal Axes) in United States and Japan.

preceded the Great Moderation. For example, in Japan prior to 1984, the relationship is linear and positive. That this result is consistent with much of the earlier literature is not surprising given that the evidence of a positive relationship comes largely from empirical studies covering the great inflation period that includes the high-inflation episodes of the 1970s and early 1980s when most industrial countries experienced double-digit inflation rates. The U-shaped relationship observed over the full sample period appears to have been driven by observations coming from the Great Moderation. Similar conclusions are drawn from analyses using several formal econometric techniques, such as semiparametric regression, rolling regression, and the multiple structural break test developed by Bai and Perron (1998, 2003). Not only does the relationship vary over time, but the time-varying pattern coincides well with changes in inflation or monetary policy regimes, possibly reflecting changes in the central bank's target for inflation or the public's inflationary expectations.[6]

Besides these empirical results, this paper presents a simple theoretical model that can accommodate the patterns observed in the data. Given that models are judged on their ability to match the data, traditionally popular models, such as menu cost or Lucas-type imperfect information models, are of reduced appeal as they typically predict a positive correlation between inflation and RPV in positive inflation environments.[7] Instead, I modify a Calvo sticky price model to incorporate multisector forward-looking components and *sectoral heterogeneity*, and show that this model accounts for the key empirical regularities found in this paper. The sectoral heterogeneity embedded in the modified Calvo sticky price model contributes to the model's relatively good fit of the observed empirical regularities. The intuition behind this is that in the presence of sectoral heterogeneity, sectors with relatively flexible prices respond more to an external shock than do sectors with relatively sticky prices. These different responses across sectors give rise to RPV. When high and persistent wage growth is expected, firms in the flexible sectors front-load prices because price adjustment is not allowed between the exogenously determined opportunities to adjust prices.

Simulations of this modified Calvo model suggest that the functional form of the relationship between inflation and RPV depends upon the degree of *price rigidity*. In a more rigid price setting environments, the relationship takes on a U-shape profile. In contrast, the U-shaped relationship disappears in more flexible pricing environments. If the degree of price stickiness varies in a systematic fashion with the average inflation rate, as has been often documented in the literature (e.g., Kiley 2000, Nakamura and Steinsson 2008b), then different degrees of price rigidity across inflation regimes could lead to differences in the underlying relationship between inflation and RPV. My empirical analysis of inflation targeting (IT) countries supports this view and shows that the relationship varies with the inflation regime. The relationship changed from a positive linear association during the high-inflation regimes that preceded the adoption of IT to a U-shaped association during the lower average inflation periods that followed.

This paper is structured as follows. After this introduction, Section 1 presents a brief description of the data used in the current study. Section 2 documents the econometric analyses performed using diverse econometric tools that yield qualitatively similar results with regard to the key conclusions of this study. Section 3 introduces a modified Calvo model to describe the important empirical regularities obtained in the

6. Fielding and Mizen (2008) also note that their results are sensitive to changes in the sample period. But instead of exploring the possibility of time-varying behavior of the relationship, they simply ascribe parameter instability to the restrictive parameterization of a quadratic model.

7. Some conventional models can account for the U-shaped relationship between inflation and RPV, but only around zero inflation. As shown later, however, in most cases the relationship is U-shaped around an inflation rate that is significantly higher than zero. The modified Calvo model introduced here can explain a U-shaped relationship around a nonzero inflation rate. See Section 3 for further discussion.

current study. This section also explores the factors behind the structural change of the relationship via simulation experiments. Section 4 concludes the paper.

1. THE DATA

My data set comprises *monthly* consumer price indices for the United States and Japan at the second level of disaggregation. As summarized in Table 1, the resulting series are available for 38 product categories in the United States beginning from 1978.M1 and for 47 categories in Japan starting from 1970.M1.[8] Price data for the United States come from the Bureau of Labor Statistics (BLS) and the Japanese CPI data are collected from the Statistical Library of the Statistics Bureau. Both headline CPI and core CPI are considered in each country in an effort to assess the role of traditionally volatile CPI items related to food and energy. Items listed in Table 1 in bold typeface represent the food- and energy-related items that are subtracted from aggregate inflation measure to calculate core inflation.

The primary measure of inflation used here is the monthly log-difference of the CPI that is computed from the seasonally adjusted price indices using the Census X12 method. RPV is then constructed by the weighted average of subaggregate inflation series using the s.d.

$$RPV_t = \sqrt{\sum_{i=1}^{N} \omega_i (\pi_{it} - \overline{\pi_t})^2},$$

where $\pi_{it} = \ln P_{it} - \ln P_{i,t-1}$, $\overline{\pi_t} = \sum_{i=1}^{N} \omega_i \pi_{it}$, ω_i denotes the fixed expenditure weight of ith product that sums to unity, and P_{it} represents the price index of i_{th} good at time t.[9]

8. Compared to the prices of specific individual commodities that have been adopted by some earlier studies in the literature (e.g., magazine prices by Cecchetti 1986, apparel and sporting goods prices by Kashyap 1995, food prices by Lach and Tsiddon 1992 and Reinsdorf 1994), disaggregated price indices contain much more information on the variability of relative prices by virtue of the wider coverage of CPI categories.

9. Given the nature of index data, the RPV measure adopted here should be read as relative *inflation* variability. Throughout the paper, however, I follow the tradition in the literature to and refer to this measure as RPV. Another common formulation for RPV is the coefficients of variation (CV) defined as $CV(\pi_t) = \frac{\text{s.d.}(\pi_t)}{|\overline{\pi_t}|}$. This measure of dispersion is not equivalent to s.d. especially when the mean level of inflation exhibits certain structural changes. In fact, while a large number of studies using variance or standard deviation as dispersion measure tend to provide evidence of a positive association between price dispersion and the inflation level, studies based on CV (e.g., Reinsdorf 1994, Silver and Ioannidis 2001) tend to find a negative relationship. Here I stick to s.d. as RPV measure for two reasons. First, the overwhelming majority of extant studies employ s.d. as the measure of RPV and hence this facilitates comparisons with previous studies. Second and more important, CV is not easily defined when inflation is close to zero or even negative, as it was in Japan during its disinflationary period.

TABLE 1

DATA DESCRIPTION

U.S. (1978.M1–2007.M9) [BLS] Item	Weight	Japan (1970.M1–2006.M7) [Statistics Bureau] Item	Weight
All items	100	All items	100
Cereals and bakery products	1.10	**Cereals**	2.19
Beef and veal	0.63	**Fish and shellfish**	2.45
Pork	0.41	**Meat**	1.98
Fish and seafood	0.34	**Dairy products**	1.09
Eggs	0.10	**Vegetables and seaweeds**	2.73
Milk	0.29	**Fruits**	1.03
Cheese and related products	0.25	**Oils, fats and seasonings**	1.01
Fresh fruits	0.49	**Cakes and candies**	2.17
Fresh vegetables	0.47	**Cooked food**	2.83
Nonalcoholic beverages	0.91	**Beverages**	1.45
Other food at home	1.74	**Alcoholic beverages**	1.36
Food away from home	5.99	Eating out	5.55
Alcoholic beverages	1.11	Rent	17.66
Shelter	32.78	Repairs and maintenance	2.72
Fuel oil and other fuels	0.34	**Electricity**	2.92
Electricity	2.75	**Gas**	1.71
Utility gas service	1.28	**Other fuel and light**	0.53
Household furnishings and operations	4.65	Household durables	1.11
Men’s apparel	0.70	Interior furnishings	0.33
Boy’s apparel	0.19	Bedding	0.29
Women’s apparel	1.35	Domestic utensils	0.71
Girls’ apparel	0.24	Domestic nondurables	0.71
Men’s footwear	0.23	Domestic services	0.30
Women’s footwear	0.36	Japanese clothing	0.17
New vehicles	4.98	Clothing	1.92
Used cars and trucks	1.72	Shirts and sweaters	0.97
Motor fuel	4.35	Underwear	0.42
Motor vehicle parts and equipment	0.37	Footwear	0.50
Medical care commodities	1.45	Other clothing	0.35
Medical care services	4.83	Services related to clothing	0.31
Sporting goods	0.67	Medicines and health fortification	1.22
Photographic equipment and supplies	0.08	Medical supplies and appliances	0.86
Toys	0.25	Medical services	2.41
Admissions	0.71	**Public transportation**	2.50
Educational books and supplies	0.20	**Private transportation**	7.78
College tuition and fees	1.52	Communication	3.64
School tuition and fees	0.41	School fees	2.73
Other goods and services	3.48	School textbooks	0.10
		Recreational durables	1.18
		Recreational goods	2.33
		Books and other reading materials	1.61
		Recreational services	5.88
		Personal care services	1.29
		Toilet articles	1.34
		Personal effects	0.73
		Cigarettes	0.63
		Other miscellaneous	1.87

NOTE: Bold-faced items represent the food- and energy-related items that are subtracted from aggregate inflation measure to calculate core inflation.

2. ECONOMETRIC ANALYSIS

Notwithstanding its merit, a visual inspection of scatterplots may be of reduced value in evaluating the underlying relationship between inflation and RPV, particularly when one suspects structural changes are present. In this section, more formal econometric tools are utilized to provide a better sense of the underlying relationship. Toward this end, three econometric methods are adopted: (i) semiparametric regression, (ii) rolling regression, and (iii) the multivariate multiple structural break test developed by Bai and Perron (1998, 2003).

2.1 Semiparametric Regression Analysis

Although the scatterplots in Figure 1 convincingly point toward a quadratic form of the RPV-inflation nexus, it is preferable to place as few restrictions on the functional form as possible especially when economic theory does not provide any concrete guidance as to the functional form. Parametric models are attractive because they can be accurately estimated and the fitted parametric models can be easily interpreted. But parametric models may present a misleading picture of the relationship if the underlying assumptions are violated. In this vein, a nonparametric approach provides clear advantages as it avoids restrictive assumptions on the functional form of regression. A nonparametric approach, however, is not without its drawbacks. Two clear drawbacks are that the results may not be easy to interpret and the approach may yield inaccurate estimates when the number of regressors is large. As a compromise, I employ a semiparametric approach that combines the attractive features of both parametric and nonparametric models, thus keeping the easy interpretability of the former while retaining some of the flexibility of the latter.[10]

Here I follow Fielding and Mizen (2008) and consider the following partially linear regression model,

$$RPV_t = X_t'\beta + g(\pi_t) + \varepsilon_t, \tag{1}$$

where X_t is a $(p + q) \times 1$ vector of the regressors that includes lagged terms of RPV and inflation, $X_t' = \{RPV_{t-1}, \ldots, RPV_{t-p}, \pi_{t-1}, \ldots, \pi_{t-q}\}$. $g(\cdot)$ is an unknown smooth differential function that captures the contemporaneous effect of inflation on RPV and it determines the underlying functional form of the relationship between inflation and RPV.[11]

10. Fully nonparametric regression estimators are more flexible and robust against functional form misspecifications, but their statistical precision decreases substantially if the regression model includes several explanatory variables. By combining the attractive features of parametric and nonparametric techniques, a semiparametric approach gets around the so-called "curse of dimensionality" problem while allowing for a flexible functional form.

11. If the true functional form is quadratic, for instance, $g(\cdot)$ will take the form of $g(\pi_t) = \alpha + \gamma_1\pi_t + \gamma_2\pi_t^2$. Details on the semiparametric approach are contained in Appendix A.

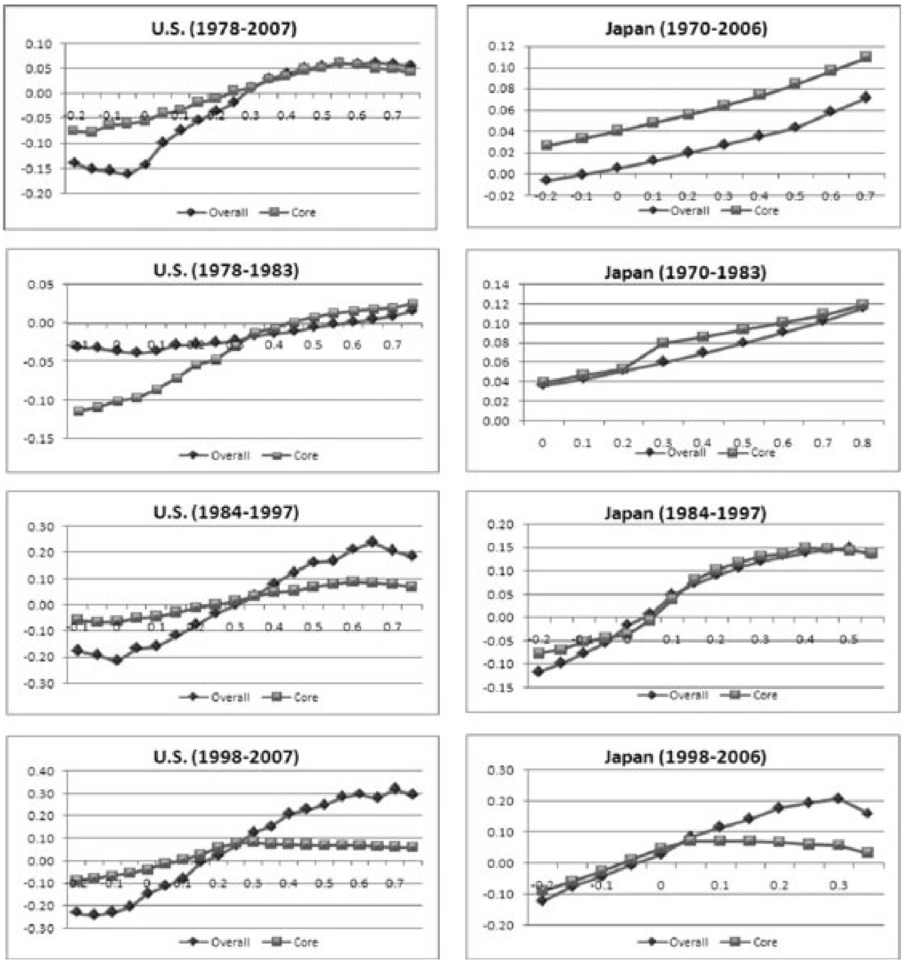

FIG. 2. Derivatives of the $g(\cdot)$ Function for Different Values of Inflation (π_t).

Figure 2 plots the semiparametric estimates of the $g'(\cdot)$ function for various values of inflation (π_t). In most cases considered, the fitted $g'(\cdot)$ function is approximately linear and upward sloping, indicating that $g(\cdot)$ is nonlinear and U-shaped. Notice the point where the estimated $g'(\cdot)$ function crosses the x-axis. At that point, $g'(\cdot) = 0$ and thus RPV is minimized at the corresponding inflation rate, which I denote as π^* throughout the paper. If the inflation rate is below π^*, then $g'(\cdot) < 0$ and $g(\cdot)$ is downward sloping, while $g'(\cdot) > 0$ and $g(\cdot)$ is upward sloping if the inflation rate is above π^*. This transition of $g'(\cdot)$ from positive to negative values indicates that $g(\cdot)$ has a quadratic form. In view of the smooth transition of the $g'(\cdot)$ function without any discontinuity around the threshold level, the underlying $g(\cdot)$ function is closer to a U-shape than to a V-shape.[12] As pointed out by Fielding and Mizen (2008), the fitted $g(\cdot)$ function is *approximately* U-shaped because the estimated $g'(\cdot)$ function is not a completely straight line but instead has a little curvature at the upper (and lower) ends.[13]

An additional advantage of a semiparametric regression method is that it enables us to track the stability of $g(\cdot)$ by examining whether the fitted $g'(\cdot)$ function shifts across samples. As shown in Figure 2, the fitted $g'(\cdot)$ function varies significantly across sample periods. In Japan, for instance, the estimated $g'(\cdot)$ function lies consistently above the x-axis for the first subsample, which suggests a monotonic relationship, while it crosses the x-axis in the subsequent subsamples. Moreover, the estimated π^* also appears to vary over subsample periods in both countries. In the United States, it shifts from 0.5% in the first subsample to around 0.1% in the third subsample, averaging around 0.25% for the full sample period. A similar shift is observed in Japan as it transitioned from trend inflation to trend deflation around the mid-1990s. The estimated π^* is around 0.05% in the second subsample but down to -0.05% in the third subsample when the Japanese economy fell into deflation. The results remain much the same when we use core inflation, which strips out the traditionally volatile prices of food and energy related items. In Japan, the exclusion of the more volatile price items does not bring about any notable difference in the results. In the United States, however, the slope of the fitted $g'(\cdot)$ function looks a bit flatter during the Great Moderation period, implying that RPV becomes less responsive to changes in the inflation level when more volatile food and energy related prices are excluded.

To sum, the results from semiparametric analysis largely confirm the visual evidence shown in Figure 1. The upward-sloping line of the fitted $g'(\cdot)$ function provides convincing evidence of U-shaped relationship and the shift across subsample periods indicates a time-varying pattern of the relationship.

2.2 Rolling Regression Analysis

So far the evidence of a time-varying pattern in the relationship between inflation and RPV has been drawn from subsample analyses in which the samples are split by cutoff points that are arbitrarily selected rather than derived. To ensure that the results are robust to the choice of cutoff points, I use a rolling regression approach that captures the time variation in the relationship without imposing any prior restriction on the timing of break points. The value of this approach is that it retains the advantage

12. Some previous studies (e.g., Tommasi 1993, Debelle and Lamont 1997) find a V-shaped relationship by regressing RPV onto the absolute values of price changes. Other studies (e.g., Parks 1978, Bomberger and Makinen 1993, Konieczny and Skrzypacz 2005) advocate employing a quadratic functional form that is compatible with the U-shaped relation witnessed here.

13. This curvature at the ends could have been driven by nonlinearity beyond the quadratic model, such as higher order polynomial terms of inflation. My simulation results posted at the author's website (http://www3.uta.edu/choi/research.htm), however, suggest that the curvature is more consistent with the quadratic model than nonlinear models containing higher order polynomial terms.

of greater flexibility in detecting structural changes over time by allowing for each rolling sample to have a completely different estimate.[14]

I estimate the following parametric model that accommodates both the inflation level (π) and squared inflation (π^2) as regressors, together with the lagged terms of RPV and inflation.

$$RPV_t = \alpha_0 + \sum_{h=1}^{p} \alpha_h RPV_{t-h} + \beta_1 \pi_t + \beta_2 \pi_t^2 + \sum_{j=1}^{q} \gamma_j \pi_{t-j} + \varepsilon_t. \tag{2}$$

Note that the inclusion of the square term of inflation is not just inspired by the pervasive evidence of a U-shaped pattern shown in Figures 1 and 2 but is also supported by some earlier studies (e.g., Parks 1978, Domberger 1987, Hartman 1991, Van Hoomissen 1988) that report inflation volatility as a significant explanatory variable of RPV.[15] Since this model specification allows the responsiveness of RPV to changes in the inflation level (π_t) and to changes in squared inflation (π_t^2) to differ, these two parameters are vital in determining the time-varying behavior of the relationship between inflation and RPV, which can be captured by their instability over rolling samples.

Before moving on, it is worth discussing briefly some important characteristics of this quadratic functional form. First, the coefficient on π_t^2, or β_2 in equation (2), captures the direction of curvature, such that the relationship takes a U-shaped function (concave upward) if β_2 is positive, while it takes an inverted U-shaped function (concave downward) if β_2 is negative. Given the compelling evidence of a U-shape, we expect the estimate of β_2 to be positive and statistically significant. If β_2 approaches zero, however, the functional form collapses to linear. In this case, the overall relationship between RPV and inflation is solely determined by β_1. Second, in the quadratic model, the minimum point of the curve provides a useful way to summarize the function. In equation (2), the minimum point pertains to the value of π at which $\widehat{RPV}$ takes on its lowest value, which corresponds to π^* as discussed in the previous section. As illustrated in Appendix B, the minimum point of the U-shaped function is reached when

$$\pi^* = \frac{-\beta_1}{2\beta_2}. \tag{3}$$

The expected sign on π^* can therefore be deduced from those of β_1 and β_2. In the U-shaped function ($\beta_2 > 0$), the sign of π^* solely depends on the sign of β_1, such that π^* is positive if $\beta_1 < 0$ but negative if $\beta_1 > 0$. As a consequence, the relationship

14. This feature of rolling regression is attractive especially when one suspects that the full-sample estimates are vulnerable to time variation in the conditional mean of the inflation process (O'Reilly and Whelan 2005).

15. These studies, however, do not present any concrete argument in relation to the U-shaped profile. Van Hoomissen (p. 1309), for example, simply states that "a quadratic term is added in the regression equation because this specification fits the data better in most cases." Domberger (p. 554) also notes that "the constrained quadratic specification is more appropriate than the unconstrained version."

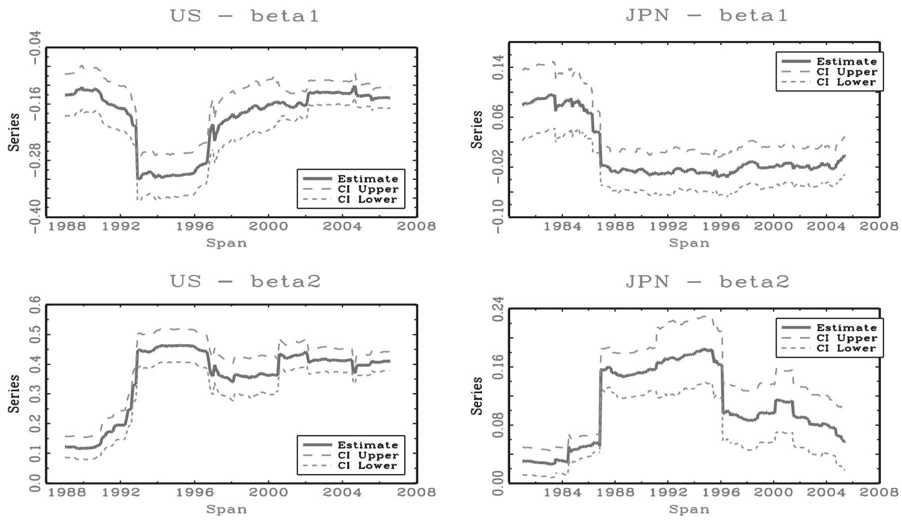

FIG. 3. Rolling 12-Year Estimates of β_1 (Top Panels) and β_2 (Bottom Panels) with the 95% Confidence Intervals.

is U-shaped around a positive inflation ($\pi^* > 0$) if $\beta_1 < 0$, but U-shaped around a negative inflation if $\beta_1 > 0$. If $\beta_1 = 0$, then $\pi^* = 0$ and the U-shape is around zero inflation. If β_2 is equal to or close to zero, however, π^* is not properly defined as it explodes. In addition, the stability of the relationship between inflation and RPV can be tracked by the time-varying behavior of π^*. As such, π^* is a crucial summary parameter in the U-shaped profile. Third, the *marginal effect* of inflation on RPV in equation (2) is not constant as in the linear model, but instead varies with inflation rates. As explained in Appendix B, the marginal effect can be approximated by

$$\frac{\Delta RPV_t}{\Delta \pi_t} \approx 2\beta_2 \pi_t + \beta_1.$$

Apparently, the change in RPV for a one-unit change in inflation depends not only on β_1 and β_2 but also on the value of π_t.[16]

Figure 3 reports the results from the rolling regression exercise by portraying the estimates of β_1 and β_2 from a sequence of rolling samples. The left- and right-hand columns of Figure 3 depict the estimates for the United States and Japan, respectively, while the top and bottom rows correspond to the estimates of β_1 and β_2. In each panel, the solid line represents coefficient estimates for β_1 (top panels) or β_2 (bottom panels) at t which are obtained using data from $t - 144$ to t with a window of 12 years,[17] while two dashed lines represent the 95% confidence interval (CI) based on

16. Notice that the marginal effect of deviation of inflation from π^*, however, solely depends on β_2 because $\frac{\Delta RPV_t}{\Delta \pi_t^d} \approx 2\beta_2 \pi_t^d$ where $\pi_t^d = \pi_t - \pi^*$ denotes the inflation deviation.

17. The choice of 12 years is mainly guided by the need for samples long enough to assess the dynamic relationship. Qualitatively similar results are obtained using rolling windows of 10 and 15 years,

the heteroskedasticity-autocorrelation robust standard errors with prewhitening. The numbers on the horizontal axis therefore represent the ending year of each 12-year window. For instance, 1990 captures the subsample period of 1979–90, and so on.

A couple of interesting features emerge from the plots. First, both $\hat{\beta}_1$ and $\hat{\beta}_2$ exhibit significant variation over time in each country, and they display similar timing of break points. Simple visual inspection shows two obvious jumps, one in 1993 and the other in 1997 for the United States and in 1987 and in 1997 for Japan. The break point of 1997 in both countries is very close to the break point of 1997 maintained in my subsample analysis of the preceding sections. It bears further noting that the width of the CI does not change in an important way. The 95% CIs of coefficient estimates are of similar magnitude across sample period and the upper and lower ends move quite closely with the coefficient estimates. Second, the estimation results are consistent with the findings from earlier sections with regard to the underlying functional form of the inflation-RPV nexus. In both countries $\hat{\beta}_2$ is consistently positive and statistically significant, indicating a U-shaped profile. The sign of $\hat{\beta}_1$, however, is monotonically negative in the United States, while it is mixed in Japan. Consequently, $\hat{\pi}^*$ as defined in (3) is consistently positive in the United States, while it swings from positive to negative in Japan. Notice that in Japan at the beginning of the sample period, which corresponds to the Great Inflation period, $\hat{\beta}_1$ is positive while $\hat{\beta}_2$ is very close to zero. This may imply that the functional form during this period is positive linear.

To further examine the evolution of π^*, I plot, in the left-hand panel of Figure 4, the rolling 12-year estimates of the annualized π^* of both countries, along with the 95% CI that is computed based on the delta method (e.g., Hayashi 2000, p. 93). Since $\hat{\pi}^*$ is not properly defined in the Great Inflation period prior to 1984 when $|\hat{\pi}^*|$ blows up as $\hat{\beta}_2$ approaches zero, I focus on the Great Moderation period from 1984 onward. The starting point on the horizontal axis therefore is 1995, corresponding to the subsample period of 1984–95. The resulting estimates of $\hat{\pi}^*$ paint a very similar picture about the time-varying feature of the relationship. $\hat{\pi}^*$ varies significantly over time, echoing the time-varying pattern of $\hat{\beta}_1$ and $\hat{\beta}_2$ in Figure 3. While $\hat{\pi}^*$ has been declining over time in both countries, it has stayed consistently above zero in the United States, but it fell from positive to negative in Japan around $2004-05$. Notice that the annualized $\hat{\pi}^*$ ranges from 2% to 4% in the United States and the lower bound of the 95% CI stays above 1.5%, pointing to a nonnegligible departure of π^* from zero. Essentially similar results are obtained from the relevant t-statistics that are computed using the asymptotic distributions of $\hat{\pi}^*$ based on the delta method. As displayed in the right-hand panel of Figure 4, the t-statistics, with the exception of Japan after 2005, are consistently greater than the 5% critical values represented by the dotted lines. That is, $\hat{\pi}^*$ is significantly different from zero by the usual 5% criteria. The statistical insignificance of $\hat{\pi}^*$ in Japan after 2005, however, does not come as a surprise in view of the fact that Japan fell into a deflationary period after 1997.

respectively, although the variability of the estimates is inversely related to the length of rolling sample window.

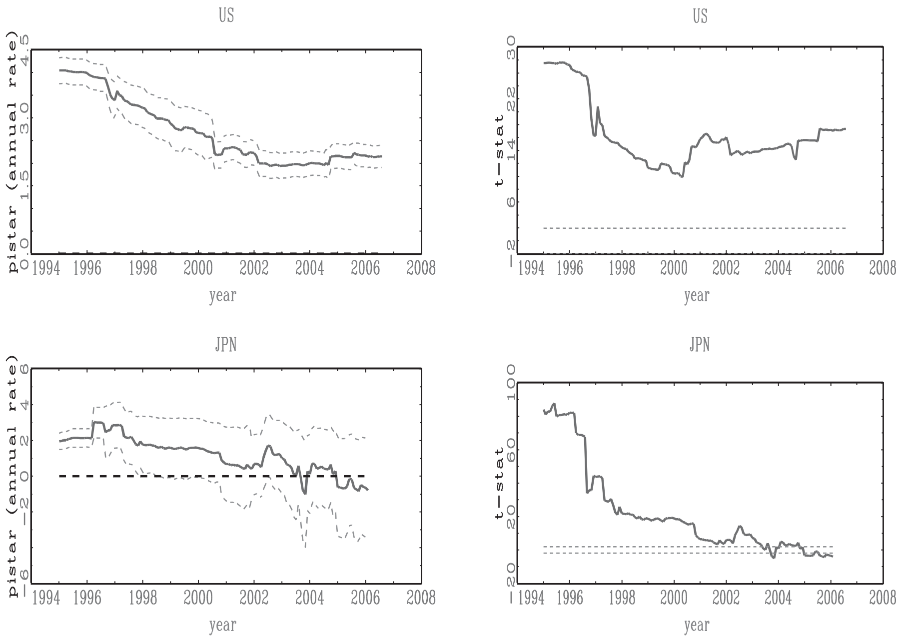

FIG. 4. Rolling 12-Year Estimates of π^* and the t-statistics.

NOTE: Panel A: $\hat{\pi}^*$ and 95% CI, Panel B: t-statistics.

Overall my results in this section strongly suggest that the relationship between inflation and RPV is U-shaped around a positive inflation rate, which is significantly different from zero. Since many existing theoretical models imply a U-shape centered at $\pi^* = 0$, my empirical results suggest that these models are inconsistent with the data and suggest the need for alternative theoretical models to match the data. This line of inquiry is pursued in Section 3.

2.3 Testing for Multiple Structural Breaks

To gain further insights into the structural change issue, I now implement the formal test for structural changes developed by Bai and Perron (1998, 2003). An appealing feature of this popular testing methodology is that it allows us to locate multiple structural breaks endogenously from the data, without assuming any prior knowledge about the potential break dates and the number of breaks. The number of breaks, their timing, and the constant are all estimated from a series of sequential Wald tests. Here I consider a multivariate setting similar to equation (2) with a general error process, in which both conditional heteroskedasticity and autocorrelation are

TABLE 2

Results of Bai–Perron Tests

Variable	Regime 1		Regime 2		Regime 3
United States					
Break point		1982:M6		1999:M9	
{95% CI}		{1982:M2,1984:M1}		{1997:M12,2000:M10}	
Constant	0.077* (0.009)		0.106* (0.007)		0.061* (0.006)
π_t	−0.062* (0.022)		−0.349* (0.025)		−0.150* (0.011)
π_t^2	0.055* (0.017)		0.480* (0.030)		0.409* (0.015)
π_{t-1}			0.004 (0.007)		
π_{t-2}			−0.000 (0.007)		
π_{t-3}			0.005 (0.007)		
π_{t-4}			−0.001 (0.007)		
π_{t-5}			−0.006 (0.006)		
RPV_{t-1}			0.123* (0.029)		
RPV_{t-2}			0.013 (0.030)		
RPV_{t-3}			−0.000 (0.030)		
RPV_{t-4}			0.057* (0.029)		
RPV_{t-5}			0.048* (0.028)		
Japan					
Break point		1986:M11		1998:M12	
{95% CI}		{1983:M10,1990:M5}		{1994:M1,2005:M10}	
Constant	0.096* (0.017)		0.058* (0.015)		0.085* (0.006)
π_t	0.090* (0.034)		−0.140* (0.028)		0.010 (0.016)
π_t^2	0.002 (0.013)		0.129* (0.008)		0.127* (0.022)
π_{t-1}			−0.003 (0.009)		
π_{t-2}			0.009 (0.008)		
π_{t-3}			0.021* (0.008)		
RPV_{t-1}			0.093* (0.041)		

Note: Maximum number of breaks set to five and minimum regime size to 5% of sample. Robust standard errors used based on a quadratic spectral kernel HAC estimator with AR(1) prewhitening filters. Standard errors are reported in parentheses. Asterisk (*) denotes significance at the 10% level.

allowed.[18] Following the guidelines from Bai and Perron, the break is assumed not to occur during the initial 15% nor the final 15% of the sample period in testing for structural breaks. The maximum number of breaks is set to five and the minimum regime size is set to 5% of the sample.

Table 2 reports the estimated dates for structural breaks in the relationship between inflation and RPV. The 95% CI is also computed using robust standard errors based on a quadratic spectral kernel HAC estimator with AR(1) prewhitening filters. Bai-Perron's multivariate break-point analysis identifies two break points (three regimes) in each country. The identified break dates are 1982 and 1999 for the United States and 1986 and 1998 for Japan. The estimated break dates are proximate to the break

18. Specifically, I employ a partial structural change model of $y_t = x_t'\beta + z_t'\delta_j + u_t$ with $t = T_{j-1} + 1, \ldots, T_j$, by setting RPV_t as y_t, $x_t = \{RPV_{t-1}, \ldots, RPV_{t-q}, \pi_{t-1}, \ldots, \pi_{t-p}\}$, and $z_t = \{c, \pi_t, \pi_t^2\}$ such that the coefficients for the inflation and squared inflation are allowed to shift. The lag lengths (p, q) in the regression equation are chosen by the AIC rule and $(p, q) = (5, 5)$ in the United States and $(p, q) = (3, 1)$ in Japan. By adding lagged terms of RPV_t as regressors, no serial correlation is assumed in the error terms, $\{u_t\}$. The reader is referred to the original work of Bai and Perron for details on the test procedures.

points maintained in my subsample analysis, 1984 and 1997, and their 95% CIs are well inclusive of the maintained break points. Table 2 also presents the parameter estimates for each regime, which are largely consistent with the results from the previous sections. In the United States, for instance, the parameter estimates for π_t and π_t^2 are statistically significant in all regimes and the parameter estimates for the two variables take opposite signs as expected, negative for β_1 and positive for β_2, indicating a U-shaped relationship around a positive level of inflation ($\hat{\pi}^* > 0$). The results in Japan take a slightly different profile. In regime 1, the coefficient estimates for both β_1 and β_2 are positive but the estimate for β_2 is neither substantially different from zero nor statistically significant, implying that the relationship in the first regime is likely to be positive linear. In regime 2, however, which covers the Great Moderation period until the onset of deflation, the estimates for both β_1 and β_2 are statistically significant. The estimate for β_1 is negative while the estimate for β_2 is positive, which indicates a U-shape profile with positive $\hat{\pi}^*$. In regime 3 for the deflationary period, the estimates for both β_1 and β_2 are positive, which suggests a U-shaped relationship around a negative $\hat{\pi}^*$, although the estimate for β_1 is statistically insignificant. Overall the results can be viewed as being supportive of the findings of previous sections as well as consistent with the maintained break points in the subsample analysis. It is reassuring to note that the three different econometric tools lead to very similar conclusions about the underlying relationship between inflation and RPV.

3. LINKING THE EMPIRICAL FINDINGS TO A THEORETICAL MODEL

The central empirical finding of this study is that the relationship between inflation and RPV not only takes a U-shaped profile, but also exhibits significant variation over time. The key aspect of this time-varying property is that it is consistent both with a U-shaped relationship and with a positive linear relationship. This empirical feature, however, is puzzling in light of the traditional workhorse models in the literature—for example, menu cost or Lucas-type imperfect information models—that predict a positive connection between inflation and RPV in a positive inflation environment (see, e.g., Fischer 1981, Lach and Tsiddon 1992, Reinsdorf 1994). As an alternative, I consider here a modified Calvo model with sectoral heterogeneity in nominal rigidity. In the presence of sectoral heterogeneity in price adjustment, RPV increases with inflation because price changes are not perfectly synchronized across sectors due to a wedge between prices with higher flexibility and prices with lower flexibility.[19] As illustrated later, the modified Calvo model fits the key stylized facts of this study quite well, especially the U-shaped profile around nonzero inflation.

19. Numerous studies provide micro- and macroevidence of significant heterogeneity in price stickiness across sectors (e.g., Blinder et al. 1999, Bils and Klenow 2004, Dhyne et al. 2006, Nakamura and Steinsson 2008a, to cite just a few). The literature provides several explanations for the heterogeneity of price rigidity across sectors, including: (i) market structure and industry concentration, (ii) contractual arrangements and the number of stages of processing, (iii) variability of the input costs, (iv) the history or persistence of local shocks and/or supply and demand elasticities, and (v) demand-side variations between durable and nondurable goods.

3.1 A Modified Calvo Model

Consider a Calvo model of sticky prices with multiple sectors, in which firms produce differentiated goods across sectors with labor as the only input to production. The production technology is linear in labor, such that

$$Y_{it} = N_{it}, \quad i = 1, \ldots, N,$$

where i denotes the ith sector.[20] In this setting, the representative firm in sector i updates its nominal prices with probability of $1 - \lambda_i$ each period, so that λ_i measures the degree of nominal rigidity in sector i. The only uncertainty in this economy is embedded in the evolution of nominal wage (W_t), which follows the stochastic process,

$$\log W_t - \log W_{t-1} = \frac{\varepsilon_t}{1 - \rho L},$$

or equivalently,

$$\log \frac{W_t}{W_{t-1}} = \rho \log \frac{W_{t-1}}{W_{t-2}} + \varepsilon_t,$$

where ε_t is white noise, L denotes the lag operator, and ρ represents the persistence of wage growth.

The sectoral price then evolves over time according to

$$\log\left(\frac{P_{it}}{W_t}\right) = \lambda_i \log\left(\frac{P_{i,t-1}}{W_{t-1}}\right) - \lambda_i \frac{(1 - \beta\rho)}{(1 - \lambda_i \beta\rho)} \log \frac{W_t}{W_{t-1}}, \tag{4}$$

where β is the usual discount factor. The detailed derivation of equation (4) is relegated to Appendix C.

3.2 Simulations

To assess the qualitative and quantitative properties of this simple forward-looking model, I run a series of simulation exercises using a range of economically meaningful parameterizations. In each simulation, I consider an environment with 30 sectors ($N = 30$) and $\beta = 0.99$, which are typical assumptions made in the literature.[21]

20. The overall economic environment of the modified Calvo sticky price model considered here is similar to that of Woodford (2003) and Yun (1996). The author is very grateful to an anonymous referee for suggesting this model. Fielding and Mizen (2008) use the Rotemberg–Danziger model with a quadratic adjustment cost function to explain the U-shaped relationship found in the U.S. data. In their model, however, since the U-shaped relationship only exists when the inflation rate is between the lower bound (π_L) and the upper bound (π_H), it is unclear whether their model can still account for the U-shaped relationship found during Japan's deflationary period, not to mention the time-varying behavior.

21. Qualitatively similar results are obtained with $N = 50$ and $N = 100$. Though attention has been restricted to a simple Calvo-type model of price stickiness, the nature of the results is likely extendable to the more general environments discussed in the recent literature (e.g., Carvalho 2006, Midrigan 2009, Nakamura and Steinsson 2008a).

TABLE 3

PARAMETER VALUES USED IN SIMULATION

	Simulation 1	Simulation 2	Simulation 3	Simulation 4
λ_i	[0.5, 0.95]	[0.5, 0.95]	[0.9, 0.95]	[0.1, 0.35]
(Duration in months)	(1.5, 19.5)	(1.5, 19.5)	(9.5, 19.5)	(0.5, 1)
β (Time-discount factor)		0.99		
ρ (Persistence of wage shock)		0.90		
σ_ε (s.d. of innovation to nominal wage)		0.01		
N (Number of sectors)		30		

NOTE: Duration of price rigidity is calculated by $d = \frac{-1}{ln(\lambda_i)}$.

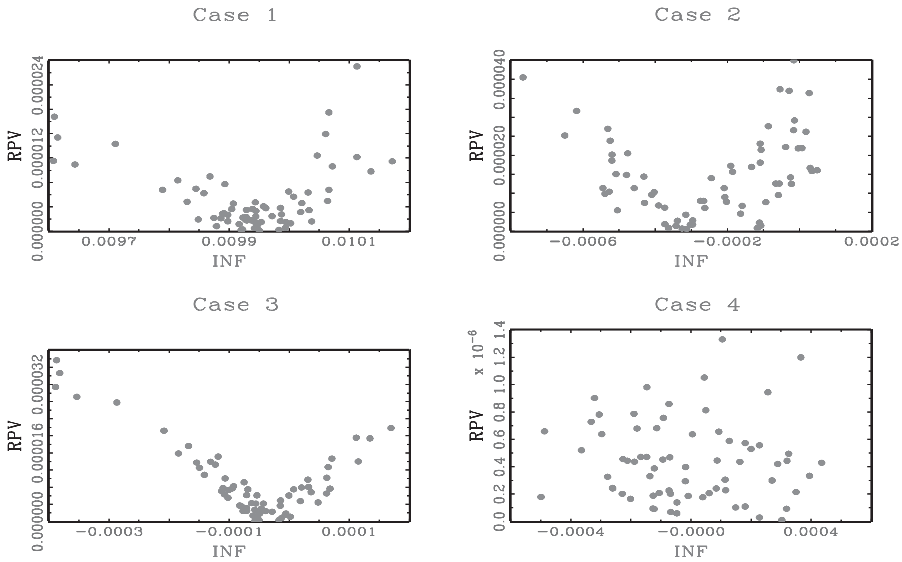

FIG. 5. Simulation Results of the Calvo Model.

The remaining structural parameter values used in the simulation experiments are specified in Table 3.

In the first two simulations, the probability that firms do not reoptimize their nominal price each period is assumed to be uniformly distributed from 0.5 to 0.9 ($\lambda_i \in \{0.5, 0.9\}$), roughly matching the 4–5 months of price stickiness of Bils and Klenow (2004).[22] The sole difference between the first two simulations lies in the inflationary environment assumed, with simulation 1 having moderate inflation and simulation 2 having moderate deflation. The top two panels of Figure 5 plot the

22. To understand this, note that the duration of price rigidity is calculated by $d = \frac{-1}{ln(\lambda_i)}$. When $\lambda_i \in \{0.5, 0.9\}$, the corresponding duration is $d \in \{1.4, 9.5\}$ months. As λ decreases, therefore, the probability that firms adjust their prices increases.

resulting simulated relationship between aggregate inflation and RPV. As it is clear from the figure, the simulated relationship is approximately U-shaped in both the mild inflation and deflation environments. The basic idea behind the U-shaped relationship generated by this simple model is that, in the presence of sectoral heterogeneity, sectors with relatively flexible prices respond more while sectors with relatively sticky prices respond less in the face of wage growth. Expectations of high and persistent wage growth lead firms in the flexible sectors to front-load prices substantively when they reset their prices. This is because, with sticky prices, a firm is not permitted to adjust its prices to account for changes in marginal costs outside of the exogenously determined opportunities.

The simulation also provides an important insight into the time-varying feature of the relationship. The bottom two panels of Figure 5 present simulated relationships assuming different levels of price stickiness. To demonstrate how the degree of price rigidity affects the relationship, λ_i is confined in simulation 3 to a narrow range of $\{0.9, 0.95\}$, which corresponds to a stickier price environment. In simulation 4, λ_i is set to $\{0.1, 0.35\}$ for a much more flexible price adjustment. As shown in the bottom-left panel, the U-shaped relationship is stronger in a more rigid price setting environment. In stark contrast, the U-shaped relationship disappears completely in the flexible price adjustment environment as displayed in the bottom-right panel. It can therefore be deduced that the degree of price rigidity exerts an important influence upon the relationship between inflation and RPV and on the relationship's time-varying behavior.

To the extent that the degree of price rigidity is systematically related to the inflation regime, as noted by Fischer (1981) and Kiley (2000),[23] we may expect changes in the degree of price rigidity driven by structural changes in the inflation process to lead to a time-varying relationship between inflation and RPV. If firms set their prices less flexibly in low-inflation settings, then the relationship will take a U-shape profile in the low-inflation environment as prices become more sticky. But the U-shaped link breaks down in the high-inflation environment when prices are adjusted more flexibly. In fact, there is a large body of evidence in the literature that uses microdata to establish a systematic link between the inflation regime and the frequency of price changes, with a higher frequency of price changes in a higher inflation environment. Cecchetti (1986) and Kashyap (1995), for example, report that the frequency of price changes in specific individual commodities, such as magazines and apparel, increases during periods of higher overall inflation. More recently, using BLS microdata, Nakamura and Steinsson (2008b) also show that the frequency of price changes is strongly related to the rate of inflation.

In light of the ample evidence the literature provides on the structural changes in the inflation process, which are driven by changes in the monetary policy regime,

23. Using a sample of 43 countries over the period of 1948–96, Kiley (2000) finds that an increase in the inflation rate leads to greater price flexibility. Devereux and Yetman (2002) consider the endogeneity of price rigidity in a Calvo model in which the degree of price rigidity (λ) is endogenously determined. Through a simple numerical approach, they find an inverse relationship between λ and average inflation rates, such that prices are less rigid when average inflation is higher.

the time-varying behavior of the relationship between inflation and RPV may result from changes in the degree of price rigidity that are caused by changes in the inflation regime. This point is partly verified by the visual evidence in Figure 1, where the relationship takes on a U-shape in the low-inflation environment after the Great Moderation. No such evidence of a U-shaped relationship is found during the high inflation of the 1970s.

3.3 Empirical Evidence from IT Countries

In order to throw additional light on the time-varying behavior of the relationship across inflation regimes, I analyze the 10 countries that have adopted IT as a framework for monetary policy. In view of the broad agreement that IT has contributed to a reduction in inflation level, the study of these IT countries serves as a useful vehicle for investigating whether the relationship between inflation and RPV varies over different inflation regimes induced by changes in the monetary policy framework. The 10 countries considered here are Australia (AUS), Brazil (BRA), Canada (CAN), Israel (ISR), Korea (KOR), Mexico (MEX), Norway (NOR), Sweden (SWE), the United Kingdom (UK), and South Africa (ZAF). Due to the limited data availability for subaggregate price indices, the sample runs from 1984 which embarks on the Great Moderation period.[24] Table 4 presents the estimation results of equation (2) for the 10 IT countries before and after the adoption of IT, together with the summary statistics on inflation.

The results in Table 4 illustrate several interesting points with regard to the relationship between inflation and RPV. First, the adoption of IT has a significant impact on average inflation. As summarized in column 2 of the table, every country under study experienced a large decline in mean inflation after IT adoption. Not surprisingly, the fall in mean inflation has been more pronounced in developing countries that started with higher inflation, such as BRA, ISR and MEX, each of which moved from double- or triple-digit annual inflation to single-digit annual inflation.

Second and more important, whether or not the decline in mean inflation induced by IT adoption brings about a change in the underlying relationship hinges upon the inflation regime prior to the adoption of IT. Note that there exists a remarkable difference in the inflation-RPV nexus between the three high-inflation countries and the others before IT adoption. Again, the point estimates of the coefficients β_1 and β_2, reported in columns 3 and 4 of Table 4, are highly informative about the functional form of the relationship. In the countries with initially low inflation rates, the relationship took on a U-shaped profile even before the adoption of IT, judging from the positive sign of $\hat{\beta}_2$ and the negative sign of $\hat{\beta}_1$. By contrast, the signs of both $\hat{\beta}_2$ and $\hat{\beta}_1$ are positive in the three high-inflation countries with the value of $\hat{\beta}_2$ being either zero or very close to zero, indicating that the underlying relationship is likely to be positive linear. This closely resembles the case of the United States and Japan in

24. Details about the data are provided in Appendix D.

TABLE 4

Estimation Results from the Selected Inflation Targeting Countries

	Inflation	$\hat{\beta}_1$ (s.e.)	$\hat{\beta}_2$ (s.e.)	$\hat{\pi}^*$ [5%, 95%]
		Before IT adoption		
AUS	5.7	−0.64 (0.09)	0.26 (0.01)	5.0‡ [3.6, 6.4]
BRA	151.4	0.07 (0.00)	0.00 (0.00)	359.8‡ [165.0, 554.5]
CAN	4.1	−0.10 (0.03)	0.26 (0.01)	2.4 [−4.3, 9.0]
ISR	35.9	0.02 (0.00)	0.01 (0.00)	−16.7 [−138.9, 105.5]
KOR	5.5	−0.56 (0.01)	0.68 (0.00)	5.0‡ [4.0, 5.9]
MEX	31.2	0.29 (0.01)	0.01 (0.00)	−283.7 [−888.0, 320.7]
NOR	3.8	−0.33 (0.02)	0.56 (0.02)	3.6‡ [2.0, 5.2]
SWE	6.2	−0.21 (0.02)	0.47 (0.00)	2.7‡ [−0.2, 5.7]
U.K.	6.0	−0.10 (0.02)	0.22 (0.00)	2.8 [−3.9, 9.6]
ZAF	11.1	−0.36 (0.03)	0.34 (0.01)	6.3‡ [2.8, 9.8]
		After IT adoption		
AUS	2.7	−0.34 (0.03)	0.19 (0.00)	3.7‡ [1.4, 5.9]
BRA	6.8	−0.10 (0.01)	0.22 (0.00)	2.8 [−2.0, 7.7]
CAN	2.1	−0.51 (0.00)	1.55 (0.01)	2.0‡ [1.5, 2.5]
ISR	2.7	−0.04 (0.01)	0.11 (0.00)	2.4 [−5.4, 10.1]
KOR	3.2	−0.25 (0.01)	0.75 (0.03)	2.0‡ [0.6, 3.4]
MEX	4.6	−1.75 (0.05)	2.36 (0.09)	4.4‡ [4.1, 4.8]
NOR	1.9	−0.30 (0.01)	0.76 (0.00)	2.4‡ [1.3, 3.5]
SWE	1.6	−0.21 (0.01)	0.66 (0.02)	1.9‡ [0.2, 3.7]
U.K.	1.9	−0.34 (0.01)	1.50 (0.05)	1.3‡ [0.8, 1.9]
ZAF	5.8	−0.88 (0.03)	0.87 (0.02)	6.1‡ [4.2, 7.1]

Note: Data description is presented in Appendix D and the regression equation is given in (2). "Inflation" represents the period average annualized monthly rates of inflation. $\hat{\beta}_1$ and $\hat{\beta}_2$, respectively, denote the effect of inflation (π) and inflation volatility (π^2) on RPV. $\hat{\pi}^*$ denotes the *annualized* monthly inflation rate at which RPV is minimized, which is computed from $\hat{\pi}^* = \frac{-\hat{\beta}_1}{2\hat{\beta}_2}$. 5% and 95% inside the squared bracket represent the lower and upper bounds of the 95% CI of $\hat{\pi}^*$ that is derived from the delta method. ‡ represents that $H_0 : \pi^* = 0$ can be rejected at 5%.

the Great Inflation period, during which the two countries experienced double-digit inflation rates.

However, no such difference can be found between high- and low-inflation countries after the adoption of IT when every country now takes a U-shaped profile following the decline of average inflation. For the three high-inflation countries, this observation suggests that the adoption of IT has led to a change in the underlying relationship, from a positive linear relationship in a high-inflation environment before IT adoption to a U-shaped relationship in a low-inflation environment after IT adoption. This mirrors similar results found in the United States and Japan before and after the Great Moderation period. Taken as a whole, the estimation results lend credence to the view that the underlying relationship differs in a systematic manner depending on the average inflation rate, nearly positive in high-inflation regime, but a U-shaped profile in low-inflation regime.

Third, the estimates of π^* also show considerable variation across samples and across countries as presented in the last column of Table 4. In the high-inflation countries, the large values of $\hat{\pi}^*$ before the adoption of IT result in large part from the small values of $\hat{\beta}_2$, probably due to the positive linear underlying relationship. In the remaining seven countries with relatively low initial inflation rates, the annualized $\hat{\pi}^*$ is positive before IT adoption, ranging from 2.4% in SWE to 6.3% in ZAF. After IT adoption, however, $\hat{\pi}^*$ is positive in all countries, at the range of $1.3\% - 6.1\%$. Apart from ISR and BRA, the lower end of the 95% CI is above zero in all countries, implying that π^* is positive and significantly different from zero. This finding can be interpreted as vindicating the U-shaped relationship around a positive π^*, which standard macroeconomic models have difficulty addressing.

Overall, the regression results based on IT countries not only accord very closely with the main findings of earlier sections but also support the main predictions of the modified Calvo model introduced in the current paper.

4. CONCLUDING REMARKS

It has long been popularly believed that RPV is positively correlated with inflation such that higher inflation levels are associated with increased cross-sectional dispersion in relative prices. The present study offered interesting new empirical and theoretical insights regarding the relationship between inflation and RPV. Using disaggregated CPI data for the United States and Japan, this study revealed that the overall relationship between inflation and RPV is nonlinear and approximately U-shaped: RPV is positively related to inflation on one side but negatively correlated on the other. This finding supports the recent empirical results of Fielding and Mizen (2008) but runs counter to most other extant findings in the literature. The disagreement between the findings in this paper and much of the previous literature, however, is not difficult to reconcile once we take into account the time-varying behavior of the relationship. Using a diverse set of econometric tools, my analysis provided compelling evidence that the relationship between inflation and RPV has not been stable over time but instead has varied significantly in a manner coinciding with changes in the monetary policy or inflation regimes. The relationship was monotonic during the high-inflation period of the 1970s and the early 1980s, as has been documented by a number of previous studies. After the Great Moderation, the relationship exhibits a U-shape profile.

In this sense, traditionally popular theoretical models, such as standard menu cost or imperfect information models, do not fit the data well as they typically predict a positive association between inflation and RPV in a positive inflation environment. As an alternative theoretical explanation, this paper presented a modified version of the Calvo sticky price model that embeds sectoral heterogeneity in price rigidity. In fact, the discussions in Bils and Klenow (2004) and Nakamura and Steinsson (2008b) suggest that the Calvo pricing scheme is generally supported by microdata on prices. This attractive feature of Calvo pricing is distinctive, especially in comparison with another popular time-dependent pricing model, the staggered price setting of Taylor (1980), which predicts a monotonic relationship between inflation and RPV. Along with the well-known appealing characteristics of the Calvo model regarding easy aggregation and simple laws of motion for price, the model's heterogeneity in the flexibility of price adjustment across sectors contributes to the relatively good fit of the empirically observed patterns. In the presence of sectoral heterogeneity, the modified Calvo model could yield not only a U-shape profile around a positive level of inflation but also time-varying behavior in the relationship.

The intuition behind this is that RPV is induced by a wedge across heterogeneous sectors in terms of the flexibility with which prices adjust. Prices are more responsive to external shocks in sectors with relatively flexible prices than in sectors with stickier prices. Insofar as the degree of price stickiness varies systematically with the inflation regime, such that it decreases with the average inflation rate, the relationship takes a U-shape in a low-inflation environment where prices are stickier. In a high-inflation environment, when price setting becomes more flexible, the U-shaped profile vanishes. To substantiate this claim, I presented empirical evidence from IT countries, in which the positive linear relationship that is found for countries that start with a high inflation rate gives way to a U-shaped profile after the adoption of IT induces a decline in the average inflation rate. Also, there is strong microdata evidence (e.g., Cecchetti 1986, Midrigan Forthcoming, Nakamura and Steinsson 2008b) that the frequency of price changes varies with the inflation regime in a systematic fashion, with a higher frequency of firms' price changes in a high-inflation environment.[25]

These findings have important policy implications. In high-inflation environments, where RPV is positively correlated with inflation, monetary authorities can improve welfare through disinflationary policies that lower RPV. In low-inflation environments, however, pursing a disinflationary policy will not necessarily lead to a welfare gain if the relationship between inflation and RPV is U-shaped. The time-varying nature of the relationship between inflation and RPV makes the central bank's task even more difficult, particularly when the degree of price rigidity changes with the inflation regime and monetary policy framework. A proper understanding of the relationship between inflation and RPV is, therefore, of great importance for policymaking, and the current paper adds to the existing literature in this regard.

APPENDIX A: PROCEDURE OF SEMI-PARAMETRIC ANALYSIS

The $g(\cdot)$ function in equation (1) is estimated semiparametrically in the following two stages. First, the parameter vector β is estimated from a regression equation of the form

$$RPV_t = \hat{X}_t'\beta + \eta_t,$$

25. Although my discussion here focused on time-dependent models mainly due to their empirical relevance, it would also be interesting to explore the potential of state-dependent pricing models that have been recently studied in the literature (e.g., Klenow and Kryvtsov 2008, Midrigan 2009, Forthcoming) to explain the empirical regularities found in this study regarding the inflation-RPV nexus. I leave further investigation of this issue to future research.

where $\hat{X}_t$ is the residual series from a nonparametric regression of X_t onto π_t. Next, the function $g(\cdot)$ is estimated nonparametrically from the regression

$$\hat{\eta}_t = g(\pi_t) + u_t,$$

where $\hat{\eta}_t = RPV_t - \hat{X}_t'\hat{\beta}$. The object of my ultimate interest is the derivative of the $g(\pi_t)$ function, $g'(\pi_t) = \frac{\partial g(\cdot)}{\partial \pi_t}$, which captures the marginal effect of π_t on $g(\cdot)$, or the sensitivity of RPV to marginal increase in inflation. If $g'(\pi_t) > 0(g'(\pi_t) < 0)$, then RPV is increasing (decreasing) with inflation. The level of inflation that ensures $g'(\pi_t) = 0$ therefore pertains to the *threshold* level of inflation at which RPV is minimized. To estimate the $g(\cdot)$ function, I adopt the Nadaraya–Watson kernel regression estimator using the Gaussian kernel based on automatic bandwidth selection method. As is widely agreed in the literature, the results of nonparametric regression are more sensitive to the bandwidth parameter than to the choice of kernel function. In my analysis, however, I find that changing the bandwidth is not much consequential to the general shape of the function.

APPENDIX B: QUADRATIC FUNCTIONAL FORM, MARGINAL EFFECTS, AND π^*

Given the nonlinear nature of the relationship, it is not straightforward to interpret the marginal effects of inflation on RPV as it depends on inflation level as shown later. To illustrate this, first let us suppose y takes a standard form of a quadratic function as

$$y = ax^2 + bx + c, \tag{B1}$$

where a is assumed to be positive to make the function as U-shape or *parabola*. This quadratic functional form can be also represented as a *vertex* form of

$$y = a(x - h)^2 + k = ax^2 - 2ahx + ah^2 + k, \tag{B2}$$

where (h, k) is the vertex of the parabola or U-shape and a determines the size and direction of the U-shape. The larger the absolute value of a, the steeper the U-shape is, and hence the value of y is increased more quickly. It is important to note that y is minimized when $x = h$. From (B1) and (B2), we obtain that $b = -2ah$ and $c = ah^2 + k$. Or equivalently,

$$h = \frac{-b}{2a},$$
$$k = c - \frac{b^2}{4a}.$$

If we replace y with RPV_t and x with π_t in (B1)

$$RPV_t = \beta_2\pi_t^2 + \beta_1\pi_t + \alpha, \tag{B3}$$

h in (B2) corresponds to $\frac{-\beta_1}{2\beta_2}$ in (B3). In other words, RPV is minimized when $\pi_t = \frac{-\beta_1}{2\beta_2}$ which I denote as π^* throughout the paper. Note that π^* is *positive* if $\beta_1 < 0$ and $\beta_2 > 0$, while π^* is negative if $\beta_1 < 0$ and $\beta_2 > 0$. In (B3), the marginal effect of inflation on RPV can be captured by

$$\frac{\Delta \mathrm{RPV_t}}{\Delta\pi_t} \approx 2\beta_2\pi_t + \beta_1,$$

which obviously depends upon the level of inflation (π_t). It should be noted that for $\pi_t < \pi^*$ an increase in inflation actually decreases RPV so long as $\beta_1 < 0$ and $\beta_2 > 0$.

APPENDIX C: DERIVATION OF THE CALVO MODEL OF STICKY PRICE

Assume that with the probability of $1 - \lambda$, firm i can change its price, P_{it}^*. Let $MC_{i,t}$ be the nominal marginal cost of production ($MC_{it} = W_t$ in my case where I assume a simple production function with labor as the only production factor and technology linear in labor) and P_t be the aggregate price index given by

$$P_t = \left[\int_0^1 P_{it}^{1-\eta} di\right]^{\frac{1}{1-\eta}}.$$

Profit maximization problem is

$$\max_{P_{it}^*} V_{it} = E_t \sum_{k=0}^{\infty} Q_{t+k|t} \left(\frac{P_t}{P_{t+k}}\right) \lambda^k \left(P_{it}^* - MC_{i,t+k}\right) \left(\frac{P_{it}^*}{P_{t+k}}\right)^{-\eta} D_{t+k},$$

where V_{it} is the nominal value of firm i; $Q_{t+k|t}$ is the *cumulative* real discount factor

$$\begin{aligned} Q_{t+k|t} &= Q_{t+k-1|t}\beta_{t+k}, \quad \text{for } k > 1, \\ Q_{t|t} &= 1, \end{aligned}$$

where $\beta_{t+k} = (1 + r_{t+k})^{-1}$ is the *period-by-period* real discount factor, P_{it}^* is the optimal choice of the price set by the firm, D_{t+k} is the aggregate factors of the economy that are composed of the sectoral total demand, economywide total demand, and aggregate price indices. We can then think of $Q_{t+k|t}(\frac{P_t}{P_{t+k}})$ as the cumulative nominal discount factor.

The first-order condition is given by

$$\frac{\partial V_{it}}{\partial P_{it}^*} = 0 = E_t \sum_{k=0}^{\infty} Q_{t+k} \left(\frac{P_t}{P_{t+k}}\right) \lambda^k \left(\frac{P_{it}^*}{P_{t+k}}\right)^{-\eta} D_{t+k} \left[1-\eta+\eta\left(\frac{MC_{i,t+k}}{P_{it}^*}\right)\right].$$

Rewriting it, we have

$$\begin{aligned} & E_t \sum_{k=0}^{\infty} Q_{t+k} \lambda^k P_{it}^{*-\eta} P_{t+k}^{\eta-1} P_t D_{t+k} \\ &= \left(\frac{\eta}{\eta-1}\right) E_t \sum_{k=0}^{\infty} Q_{t+k} \lambda^k P_{it}^{*\eta-1} P_{t+k}^{\eta-1} P_t MC_{i,t+k} D_{t+k}. \end{aligned} \tag{C1}$$

As $1-\lambda$ fraction of firms can change prices and select price P_{it}^*, the price index follows

$$P_t^{1-\eta} = \lambda P_{t-1}^{1-\eta} + (1-\lambda) P_{it}^{*1-\eta}. \tag{C2}$$

Let us assume that the inflation rate is zero in the steady state. With this assumption, we have $P = P_i^* = (\frac{\eta}{\eta-1})MC$ in the steady state.

Now, we use the log-linearization to (C1) and (C2). From (C1), we have

$$p_{it}^* = (1-\beta\lambda) \sum_{j=0}^{\infty} (\beta\lambda)^j E_t mc_{t+j}, \tag{C3}$$

where $x = (X_t - X)/X$. From (C2), we have

$$p_t = \lambda p_{t-1} + (1-\lambda) p_{it}^*. \tag{C4}$$

Now, we assume $MC_t = W_t$, and W_t follows

$$\log \frac{W_t}{W_{t-1}} = \rho \left(\log \frac{W_{t-1}}{W_{t-2}}\right) + u_t.$$

We can rewrite it as

$$w_{t+j} = w_{t+j-1} + \frac{u_{t+j}}{1-\rho L},$$

where L is the lag operator. With repeated iteration, we have

$$w_{t+j} = w_t + \left(\frac{1}{1-\rho L}\right)\left(u_{t+j} + u_{t+j-1} + \cdots + u_{t+1}\right).$$

Since $\Delta w_{t+j} \equiv w_{t+j} - w_{t+j-1} = u_{t+j}/(1-\rho L)$, and $E_t \Delta w_{t+j} = \rho^j \Delta w_t$, we have

$$\begin{aligned} E_t w_{t+j} &= w_t + \rho\left(1 + \rho + \cdots + \rho^{j-1}\right)\Delta w_t \\ &= w_t + \frac{\rho\left(1-\rho^j\right)}{1-\rho}\Delta w_t. \end{aligned} \tag{C5}$$

Applying (C5) to (C3) with $mc_t = w_t$, we have

$$\begin{aligned} p_{it}^* &= (1-\beta\lambda)\sum_{j=0}^{\infty}(\beta\lambda)^j\left[w_t + \frac{\rho\left(1-\rho^j\right)}{1-\rho}\Delta w_t\right] \\ &= (1-\beta\lambda)\sum_{j=0}^{\infty}(\beta\lambda)^j\, w_t + \rho\left(\frac{1-\beta\lambda}{1-\rho}\right)\sum_{j=0}^{\infty}(\beta\lambda)^j\left(1-\rho^j\right)\Delta w_t \\ &= w_t + \rho\left(\frac{1-\beta\lambda}{1-\rho}\right)\left[\left(\frac{1}{1-\beta\lambda}\right) - \left(\frac{1}{1-\beta\lambda\rho}\right)\right]\Delta w_t \\ &= w_t + \frac{\beta\lambda\rho}{1-\beta\lambda\rho}\Delta w_t. \end{aligned} \tag{C6}$$

From (C4) and (C6), we have

$$p_t = \lambda p_{t-1} + (1-\lambda)\left[w_t + \frac{\beta\lambda\rho}{1-\beta\lambda\rho}\Delta w_t\right].$$

Rearranging it, we have

$$\begin{aligned} p_t - w_t &= \lambda\left(p_{t-1} - w_{t-1}\right) - \lambda\Delta w_t + \frac{\beta\lambda\rho\,(1-\lambda)}{1-\beta\lambda\rho}\Delta w_t, \\ &= \lambda\left(p_{t-1} - w_{t-1}\right) - \frac{\lambda\,(1-\beta\rho)}{1-\beta\lambda\rho}\Delta w_t. \end{aligned}$$

In level terms, we have

$$\log\frac{P_t}{W_t} = \lambda\log\left(\frac{P_{t-1}}{W_{t-1}}\right) - \frac{\lambda\,(1-\beta\rho)}{1-\beta\lambda\rho}\log\left(\frac{W_t}{W_{t-1}}\right).$$

APPENDIX D: Data Description of 10 Inflation Targeting Countries

Country	Data span	Data source	Subaggregate Items (weight in percent)
AUS	1984:Q1–2008:Q4 {1993:Q2}	Australian Bureau of Statistics (ABS)	[8] food (20.7); alcohol & tobacco (9.2); clothing & footwear (5.2); housing (26.2); household contents & services (12.9); transportation (17.6); communication (4.4); education (3.6)
BRA	1984:M1–2008:M12 {1999:M6}	Brazilian Institute of Geography and Statistics (IBGE)	[7] food products & beverages (23.6); housing (16.7); household articles (6.4); apparel (8.4); transportation & communication (20.1); health & personal care (10.3); personal expenses (14.6)
CAN	1984:M1–2009:M2 {1991:M2}	Statistics Canada	[8] food (17.0); shelter (26.6); household operations & furnishings (11.1); clothing & footwear (5.4); transportation (19.9); health & personal care (4.7); recreation, education & reading (12.2); alcoholic beverages & tobacco products (3.1)
ISR	1984:M1–2009:M2 {1997:M6}	Central Bureau of Statistics	[10] food, excluding vegetables and fruit (14.8); vegetables & fruit (3.6); housing (20.7); dwellings maintenance (10.6); furniture & household equipment (3.8); clothing & footwear (3.2); health (5.2); education, culture & entertainment (12.5); transport & communication (21.1); miscellaneous (4.5)
KOR	1985:M1–2009:M2 {1998:M4}	Korea National Statistical office (NSO)	[12] food & non-alcoholic beverages (14.0); alcoholic beverages & cigarettes (1.5); clothing & footwear (5.8); housing, water & fuels (17.0); furnishings & household equipment (4.2); health (5.2); transportation (10.9); communication (6.0); culture & Recreation (5.6); education (11.1); eating-out & accommodation (13.3); miscellaneous (5.4)
MEX	1984:M1–2009:M2 {2001:M1}	Bank of Mexico	[10] food, beverages & tobacco (22.7); clothes, footwear & accessories (5.6); housing (26.4); furniture & domestic accessories (4.9); health & personal care (8.6); transportation (13.4); education & entertainment (11.5); miscellaneous (6.9)
NOR	1984:M1–2009:M2 {2001:M3}	Statistics Norway	[12] food & nonalcoholic beverages (11.2); alcoholic beverages & tobacco (2.7); clothing & footwear (5.9); housing, water, electricity, gas & other fuels (29.5); furnishings, household equipment & routine maintenance (6.3); health (2.7); transport (17.9); communications (2.1); recreation & culture (12.0); education (0.3); restaurants & hotels (3.4); miscellaneous goods & services (6.0)
SWE	1984:M1–2009:M2 {1993:M1}	Statistics Sweden	[11] food & nonalcoholic beverages (13.2); alcoholic beverages & tobacco (3.7); clothing & footwear (5.4); housing, water, electricity, gas & other fuels (26.7); furnishings & household goods (5.5); health (3.2); transport (14.6); communication (3.5); recreation & culture (11.9); restaurants & hotels (6.8); miscellaneous goods & services (5.4)
U.K.	1988:M1–2009:M2 {1992:M10}	National Statistics	[12] food & nonalcoholic beverages (11.8); clothing & footwear (5.7); alcoholic beverages,tobacco & narcotics (4.4); housing, water & fuels (12.6); furnishings, household equipment & routine repair of house (6.6); health (2.2); transport (15.1); communication (2.3); recreation & culture (14.5); education (2.1); hotels, cafes & restaurants (12.8); miscellaneous goods & services (9.9)
ZAF	1984:M1–2008:M12 {2000:M2}	Statistics South Africa	[17] food (21.0); nonalcoholic beverages (1.1); alcoholic beverages (1.4); cigarettes, cigars & tobacco (1.1); clothing & footwear (3.3); housing (22.1); fuel & power (3.5); furniture & equipment (2.5); household operation (4.8); medical care & health expenses (7.2); transport (14.8); communication (3.0); recreation & entertainment (3.3); reading matter (0.4); education (3.5); personal care (3.7); other (3.3)

NOTE: Dates in the curved brackets represent the official adoption dates of inflation targeting that are used as the break point in my analysis. Numbers in the straight brackets represent the number of subaggregates, and the entries inside the parentheses denote the weight of each subaggregate.

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