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Abstract

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Exchange Rate Pass-Through and Inflation: A Nonlinear Time Series Analysis*

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Abstract

This paper investigates the relationship between the exchange rate pass-through (ERPT) and inflation by estimating a nonlinear time series model. Based on a simple theoretical model of ERPT determination, we show that the dynamics of ERPT can be well approximated by a class of smooth transition autoregressive (STAR) models using the past inflation rate as a transition variable. We employ several U-shaped transition functions in the estimation of the time-varying ERPT to US domestic prices. The estimation result suggests that declines in the ERPT during the 1980s and 1990s are associated with lowered inflation.

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1 Introduction

Within the framework of new open economy macroeconomic models, the degree of exchange rate pass-through (ERPT) into domestic prices is one of the key elements in evaluating international spill-over effects of monetary policy. Over the past decade, a number of empirical studies have investigated whether ERPT, defined as the response of domestic inflation rates to the changes in exchange rates (or in marginal costs), decreased during the 1980s and 1990s.[1] If there was a reduction in ERPT, it is natural to conjecture some interaction between the ERPT and the inflation rate because the timing corresponds, in many countries, to a period of low and stable inflation. This view is emphasized by Taylor (2000), who states that "the lower pass-through should not be taken as exogenous to the inflationary environment (p.1390)."

The purpose of this paper is to investigate Taylor's hypothesis on the positive relationship between the ERPT and inflation by estimating a nonlinear time series model. In particular, we employ the class of smooth transition autoregressive (STAR) models so that the degree of ERPT to domestic prices can be determined by the lagged domestic inflation rate. Most previous empirical studies on the positive association between ERPT and inflation focus on the cross-country evidence, including the analyses by Calvo and Reinhart (2002), Choudhri and Hakura (2006), and Devereux and Yetman (2010). This paper differs from the existing studies in that we examine the role of inflation in the time-varying ERPT under the time series modeling framework.

In the empirical literature on the nonlinear adjustment of real exchange rates, STAR models have been popularly employed in many studies, including Michael et al. (1997), Taylor and Peel (2000), Taylor et al. (2001), and Kilian and Taylor (2003), among others. However, STAR models have rarely been used in analyses

[1] See, for example, Goldberg and Knetter (1997), Otani et al. (2003), Campa and Goldberg (2005), Sekine (2006) and McCarthy (2007).

of the ERPT.[2] While nonlinear mean reversion of real exchange rates implies the full ERPT in the long-run, it does not imply the time-varying ERPT. We employ several U-shaped transition functions in STAR models to consider alternative forms of time-varying ERPT. Our method is applied to monthly US import and domestic price data and evaluates fluctuations of ERPT during the period of 1975 to 2007.

To motivate our nonlinear regression approach, we first present a simple theoretical model of importing firms where the ERPT becomes a nonlinear function of the past inflation rate. Our model is closely related to a model of ERPT developed by Devereux and Yetman (2010) so that the optimal price level depends directly on the nominal exchange rate, which corresponds to the marginal cost, and that importing firms endogenously select the probability of adjusting their price to an optimal level. However, our model differs from their model in several aspects. First, for every period, a fraction of firms make a finite-period Taylor (1980) type staggered contract of an inflation indexation rule. Second, each firm faces the problem of opting out of a contract. When firms opt out, they can set an optimal price by paying a fixed cost. Because the ERPT increases if more firms set an optimal price, and the probability of opting out depends on the past inflation rate, our model predicts that ERPT depends on the lagged inflation. This prediction is in contrast to the case of Devereux and Yetman (2010) where the ERPT depends on the steady-state inflation level of the economy. We show that the dynamics of ERPT predicted by the theoretical model can be well approximated by the STAR structure, and that the past decline during the 1980s and 1990s and the recent increase in the ERPT to US prices are well explained by the STAR model.

The remainder of the paper is organized as follows. Section 2 briefly describes the prediction from the theoretical model. Section 3 introduces the empirical model. Estimation results are provided in Section 4, followed by conclusions in Section 5.

[2] One of the few exceptions is a study of UK import prices by Herzberg et al. (2003). However, their study did not find supporting evidence on nonlinearity.

2 Theoretical Motivation

In this section, we briefly describe our theoretical model of importing firms, which predicts that the ERPT depends on the lagged inflation.[3] The basic setup is similar to Devereux and Yetman (2010) in that importing firms are monopolistic competitors who import differentiated intermediate goods from abroad. A representative domestic final good producer purchases all the imported intermediate goods and combines them to produce a final output. Pricing contracts between the importers and the final good producer are valid for $N(\geq 2)$ periods long, and a constant fraction $1/N$ of all importing firms write their contracts in any given time period. However, firms are allowed to opt out during the contract period and to deviate from the contract pricing rule by paying a fixed cost F. During the first $N^*(\geq 1)$ periods of the contract, firms follow the contract pricing rule and fully index their prices to aggregate inflation π_t of the initial contract period. If firms opt out of the contract after N^* periods, for the remaining periods of the contract $N - N^*$, they can charge the desired price $\widehat{p}_t = s_t + p_t^* + \mu$ where s_t is the nominal exchange rate, p_t^* is the foreign currency price and μ is a markup. Since the marginal cost $s_t + p_t^*$ is assumed to follow a random walk process (with the variance of its increment σ^2), all the firms entering into new contracts at time t set their price at $\widehat{p}_t$. Therefore, for firms that write their contracts at time t and opt out at time $t + N^*$, the entire price path is given by $\left\{\widehat{p}_t, \widehat{p}_t + \pi_t, ..., \widehat{p}_t + (N^* - 1)\pi_t, \widehat{p}_{t+N^*}, ..., \widehat{p}_{t+(N-1)}\right\}$.

We follow Ball et al. (1988), Romer (1990), and Devereux and Yetman (2002, 2010), among others, and re-formulate the firm's optimization behavior so that the probability of (not) changing its price to the desired price level is endogenously determined. Let $\kappa^{(t)}$ be the (conditional) probability that a firm under contract in the current period will maintain the contract price in the next period. Here, a superscript t in parenthesis signifies that this probability applies to all the firms entering into

[3] The details of the model are provided in the Appendix.

new contracts at time t, but not to firms in other cohorts. After setting the new contract price at t, the firms observe the aggregate inflation π_t and choose $\kappa^{(t)}$ to minimize the expected loss function given by

$$L_t = E_t \left[\sum_{j=1}^{N-1} (\beta\kappa^{(t)})^j (\widehat{p}_t + j\pi_t - \widehat{p}_{t+j})^2 \right] + \frac{1-\kappa^{(t)}}{\kappa^{(t)}} \sum_{j=1}^{N-1} (\beta\kappa^{(t)})^j \left(\sum_{\ell=1}^{N-j} \beta^{\ell-1} \right) F \quad (1)$$

where β is a discount factor. The above function implies that the loss is an increasing function of the inflation rate in absolute value. As the inflation rate rises (relative to the size of the fixed cost), the firm can minimize loss by avoiding the inflation indexation. This strategy leads to a lower $\kappa^{(t)}$ (or a shorter average length of N^*). In an extreme case of a high inflation, $\kappa^{(t)} = 0$ (or $N^* = 1$) is selected with a pricing path given by $\left\{\widehat{p}_t, \widehat{p}_{t+1}, ..., \widehat{p}_{t+(N-1)}\right\}$. In the other extreme case of a low inflation, $\kappa^{(t)} = 1$ (or $N^* = N$) can be selected with a pricing path given by $\{\widehat{p}_t, \widehat{p}_t + \pi_t, \widehat{p}_t + 2\pi_t, ..., \widehat{p}_t + (N-1)\pi_t\}$. In general, between the two extreme cases, the solution becomes a function of the inflation rate and can be expressed as $\kappa^{(t)} = \kappa(\pi_t)$.

The (short-run) ERPT is defined as the first derivative of π_t with respect to a change in marginal cost, $\Delta(s_t + p_t^*)$. Using the dynamic Phillips curve derived from the model, the ERPT can be expressed in terms of $\kappa^{(t-j)} = \kappa(\pi_{t-j})$ for $j = 1, ..., N-1$, so that the ERPT depends directly on the lagged inflation. When $N = 2$, the model reduces to the two-period Taylor (1980) model with a possibility of opting out in the second period as considered by Ball and Mankiw (1994) and Devereux and Siu (2007). In this simple case, the inflation dynamics follow a nonlinear AR(2) model with the ERPT given by $1 - \kappa(\pi_{t-1})/2$ where $\kappa(\pi_{t-1}) = \mathbf{1}\{|\pi_{t-1}| \leq \sqrt{F - \sigma^2}\}$. Figure 1 shows the predicted relationship between the lagged inflation rate and the ERPT. Abrupt transitions at the threshold values $\sqrt{F - \sigma^2}$ and $-\sqrt{F - \sigma^2}$ suggest the possibility of approximating the ERPT by a variant of a threshold autoregressive model (TAR), which is sometimes referred to as the three-regime TAR model or the band TAR model. When N becomes greater than 2, transitions become smoother. For example, when $N = 3$, inflation follows a nonlinear AR(3) model with the ERPT given by $1 - \{\kappa(\pi_{t-1}) + \kappa(\pi_{t-2})^2\}/3$ where

$$\kappa(\pi_t) = \frac{-(F - \sigma^2 - \pi_t^2)}{2\beta(F - 2\sigma^2 - 4\pi_t^2)},$$

provided $F - \sigma^2 - \pi_t^2 > 0$ and $(F - \sigma^2 - \pi_t^2) + 2\beta(F - 2\sigma^2 - 4\pi_t^2) < 0$. As shown in Figure 2 (which imposes $\pi_{t-1} = \pi_{t-2}$), the smooth nonlinear relationship between the inflation and the ERPT resembles the adjustment dynamics described by a class of STAR models with a U-shaped transition function using the lagged inflation rate as a transition variable.

3 Econometric Procedures

This section introduces the nonlinear time series model that we will use in the empirical analysis. There are three main predictions of the theoretical model on the ERPT we wish to incorporate in the empirical model. First, higher inflation (in absolute value) results in a higher degree of the ERPT. Second, the ERPT may be expressed as a symmetric function of the past inflation rates around zero. Finally, in general, the dynamics of the ERPT can be described as a smooth rather than an abrupt transition using the past inflation rate as the transition variable possibly with multiple lags. The only exception is a special case of two-period contract that predicts a discrete transition typically assumed in the TAR model.

To incorporate these features in a parsimonious parametric model, we primarily employ the exponential STAR (ESTAR) model, where a symmetric U-shaped transition function is represented by an exponential function

$$G(z_t; \gamma) = 1 - \exp\{-\gamma z_t^2\},$$

where z_t is a transition variable and $\gamma(> 0)$ is a parameter defining the smoothness of the transition. It is a popularly used STAR model originally proposed by Haggan and Ozaki (1981) and later generalized by Granger and Teräsvirta (1993) and Teräsvirta

(1994) among others. Since our objective is to determine the relationship between π_t and $\Delta(s_t + p_t^*)$, we estimate a bivariate variant of the ESTAR models specified as

$$\begin{aligned}\pi_t &= \phi_0 + \sum_{j=1}^{N} \phi_{1,j}\pi_{t-j} + \sum_{j=0}^{N-1} \phi_{2,j}\Delta(s_{t-j} + p_{t-j}^*) \\ &\quad + \left(\sum_{j=1}^{N} \phi_{3,j}\pi_{t-j} + \sum_{j=0}^{N-1} \phi_{4,j}\Delta(s_{t-j} + p_{t-j}^*)\right) G(z_t;\gamma) + \varepsilon_t, \qquad (2)\end{aligned}$$

where $\varepsilon_t \sim$i.i.d.$(0, \sigma_\varepsilon^2)$. Note that the lag length of π_t and $\Delta(s_t + p_t^*)$ on the right-hand side of (2) comes from the prediction of the theoretical model provided in the Appendix. While our theoretical model also suggests multiple transition variables, here we consider a parsimonious specification and use a moving average of the past inflation rates as a single transition variable, $z_t = d^{-1}\sum_{j=1}^{d} \pi_{t-j}$.[4] In this ESTAR framework, our interest is to obtain the time-varying ERPT defined as

$$ERPT = \phi_{2,0} + \phi_{4,0}G(z_t;\gamma).$$

We impose a restriction $0 \leq \phi_{2,0} \leq 1$ and $\phi_{2,0} + \phi_{4,0} = 1$ so that the ERPT falls in the range of $[0, 1]$.

In addition to the ESTAR model, our primary model in the analysis, we also consider another class of STAR models based on a different U-shaped transition function constructed from a combination of two logistic functions. This variant of logistic STAR (LSTAR) models has been considered in Granger and Teräsvirta (1993) and Bec et al. (2004) and is sometimes referred to as the three-regime LSTAR model. Here, we simply call the model a dual (or double) LSTAR (DLSTAR) model to emphasize the presence of two logistic functions.[5] The transition function in the DLSTAR model is given by

$$G(z_t;\gamma_1,\gamma_2,c_1,c_2) = (1+\exp\{-\gamma_1(z_t - c_1)\})^{-1} + (1+\exp\{\gamma_2(z_t + c_2)\})^{-1}$$

[4] As in Kilian and Taylor (2003), we can also employ the transition variable, $z_t = \sqrt{d^{-1}\sum_{j=1}^{d} \pi_{t-j}^2}$, which yields a similar parsimonious specification. The main result turns out to be unaffected even if our transition variable is replaced by this alternative.

[5] We use this terminology since the model differs from the multiple regime STAR models defined in van Dijk et al. (2002).

where γ_1, $\gamma_2(>0)$ are parameters defining the smoothness of the transition in the positive and negative regions, respectively, and $c_1, c_2(>0)$ are location parameters. The definitions of all other variables and parameters remain the same as in the ESTAR model. The function of our interest, the ERPT, is similarly computed as

$$ERPT = \phi_{2,0} + \phi_{4,0} G(z_t; \gamma_1, \gamma_2, c_1, c_2).$$

The reason for considering this alternative specification of the transition function is two-fold. First, as pointed out by van Dijk et al. (2002), the transition function in the ESTAR model collapses to a constant when γ approaches infinity. Thus the model does not nest the TAR model with an abrupt transition as predicted by the theory when there are only two cohorts of firms in the economy. In contrast, the DLSTAR model includes the TAR model by letting γ_1 and γ_2 tend to infinity. Second, and more importantly, the model can incorporate both symmetric ($\gamma_1 = \gamma_2 = \gamma$ and $c_1 = c_2 = c$) and asymmetric ($\gamma_1 \neq \gamma_2$ and $c_1 \neq c_2$) adjustments between the positive and negative regions. Therefore, we can investigate the case beyond our simple model that predicts a symmetric relationship between the ERPT and the lagged inflation rate. In the estimation of DLSTAR models, we employ both specifications of symmetric and asymmetric adjustments.

Note that all specifications in our analysis can be represented as

$$\pi_t = x_t' \phi_1 + G(z_t; \theta) x_t' \phi_2 + \varepsilon_t,$$

where $x_t = (1, \pi_{t-1}, ..., \pi_{t-N}, \Delta(s_t + p_t^*), ..., \Delta(s_{t-N+1} + p_{t-N+1}^*))'$, $z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$ and $\theta = \gamma$ for ESTAR models, $\theta = (\gamma, c)'$ for symmetric DLSTAR models, $\theta = (\gamma_1, \gamma_2, c_1, c_2)'$ for asymmetric DLSTAR models, respectively. In our analysis, we follow van Dijk et al. (2002) and employ the Lagrange multiplier (LM)-type test for linearity against the class of STAR models, based on the artificial model of the form:

$$\pi_t = x_t' \beta_0 + x_t' z_t \beta_1 + x_t' z_t^2 \beta_2 + x_t' z_t^3 \beta_3 + e_t. \tag{3}$$

Let $\widetilde{e}_t = \pi_t - x_t'\widetilde{\beta}_0$ be the regression residual from (3) with restrictions $\beta_1 = \beta_2 = \beta_3 = 0$ and $\widehat{e}_t$ be the residual from the full regression (3). Then, the LM test statistic can be computed as $LM = T(SSR_0 - SSR_1)/SSR_0$ where $SSR_0 = \sum \widetilde{e}_t^2$ and $SSR_1 = \sum \widehat{e}_t^2$. The LM statistic asymptotically follows χ^2 distribution with $3(2N+1)$ degree of freedom under the null hypothesis of linearity. To improve the finite sample size property, Teräsvirta (1994) also recommends the F version of the LM test statistics given by

$$F_L = \frac{(SSR_0 - SSR_1)/3(2N+1)}{SSR_1/(T - 4(2N+1))}.$$

The F statistic approximately follows F distribution with $3(2N+1)$ and $T - 4(2N+1)$ degrees of freedom under the null hypothesis. In addition, we also employ a heteroskedasticity-robust variant of the LM test suggested by Granger and Teräsvirta (1993) and denote the test statistic by LM^*.

As discussed in Teräsvirta (1994), the auxiliary regression (3) can be further used to choose the specification among alternative STAR models. In our context, the F test for $H_0 : \beta_3 = 0$ against $H_1 : \beta_3 \neq 0$ can be used as a test for an ESTAR model against an asymmetric DLSTAR model (F_3). Similarly, the F test for $H_0 : \beta_1 = 0|\beta_3 = 0$ against $H_1 : \beta_1 \neq 0|\beta_3 = 0$ can be used as a test for a symmetric DLSTAR model against an ESTAR model ($F_{1|3}$). Finally, the F test for $H_0 : \beta_1 = \beta_3 = 0$ against $H_1 : \beta_1 \neq 0$ and $\beta_3 \neq 0$ can be used as a test for a symmetric DLSTAR model against an asymmetric DLSTAR model (F_{13}).[6]

[6] The set of restrictions follows from the fact that a third-order Taylor approximation of the transition function of a symmetric DLSTAR model is given by $G(z_t; \gamma, c) \approx \left(\gamma^3 c/24 - \gamma c/2\right) + (\gamma^3 c/8) z_t^2$, while both z_t and z_t^3 appear for an asymmetric DLSTAR model similar to a single LSTAR model. The details of the derivation are available from the authors upon request.

4 Empirical Results

4.1 Data and the Linearity Test

All the data we use in the STAR estimation are taken from *International Financial Statistics* (IFS) of the International Monetary Fund. First, the main regressor in the ERPT regression is the monthly log changes in nominal exchange rate and import price in foreign currency. Since the Bureau of Labor Statistics constructs the US import price index using US dollar prices paid by the US importer, $\Delta(s_t + p_t^*)$ is simply computed as $100 \times (\ln IMP_t - \ln IMP_{t-1})$ where IMP_t is the import price after making a seasonal adjustment using X-12-ARIMA procedure. The import prices are based either on "free on board (f.o.b.)" foreign port or "cost, insurance, and freight (c.i.f.)" US port transaction prices, depending on the practices of the individual industry. In either case, under our assumption of a constant iceberg transaction cost (proportional to import price in domestic currency), the same formula can be used to compute the monthly log changes in the prices of imported goods, excluding the cost of transaction. Second, for the inflation used for the dependent and transition variables, we employ the producer price index rather than the consumer price index since the domestic price in our model is the price at which the final good producer sells its product.[7] The monthly log inflation π_t is computed as $100 \times (\ln PPI_t - \ln PPI_{t-1})$ where PPI_t is the seasonally adjusted US producer price index. As shown in Figure 3, our sample period from January 1975 to December 2007 covers the high inflation episodes in the late 1970s and the relatively stable inflation environment beginning in the 1980s, as well as the recent resurgence of a hike in the oil prices.

As a preliminary test, we first conduct the LM tests of linearity against the STAR alternatives. The results using $N = 6$ and d between 1 and 6 are reported in Table 1. All three tests, LM, F_L and LM^*, strongly suggest the presence of nonlinearity in inflation dynamics for all values of d.

[7] There are other studies that also report ERPT to producer price index. See, for example, Choudhri et al. (2005) and McCarthy (2007).

4.2 ESTAR Model

For the estimation of the ESTAR model, our primary model in the analysis, we first search for the length of moving average d in the transition variable z_t that best fits the specification. We fix the lag length $N = 6$ and search for the value of d between 1 and 6 that minimizes the residual sum of squares from the nonlinear least squares regression of (2). This search procedure leads to the choice of $d = 3$. We then adopt a general-to-specific approach, as suggested by van Dijk et al. (2002), in arriving at the final specification. Starting with a model with $N = 6$, we sequentially remove the lagged variables for which the t statistic of the corresponding parameter is less than 1.0 in absolute value. The resulting final specification and the estimates for the ESTAR model are as follows:

$$\begin{aligned}
\pi_t \;=\;& \underset{(3.118)}{0.099} + \underset{(2.322)}{0.123}\pi_{t-1} + \underset{(4.706)}{0.200}\pi_{t-3} - \underset{(1.689)}{0.081}\pi_{t-4} + \underset{(9.746)}{0.336}\Delta(s_t + p_t^*) \\
& + \underset{(2.803)}{0.093}\Delta(s_{t-1} + p_{t-1}^*) + \underset{(1.859)}{0.074}\Delta(s_{t-4} + p_{t-4}^*) + \underset{(1.349)}{0.039}\Delta(s_{t-5} + p_{t-5}^*) \\
& + \Bigg[\underset{(2.103)}{0.752} - \underset{(3.400)}{1.352}\pi_{t-5} + \underset{(19.246)}{0.664}\,\Delta(s_t + p_t^*) - \underset{(2.849)}{0.569}\Delta(s_{t-2} + p_{t-2}^*) \\
& - \underset{(1.393)}{0.300}\Delta(s_{t-4} + p_{t-4}^*)\Bigg] G\left(z_t; \widehat{\gamma}\right) + \widehat{\varepsilon}_t,
\end{aligned}$$

$$G\left(z_t; \widehat{\gamma}\right) = 1 - \exp\left\{ \underset{(4.777)}{-0.076} \left(\frac{1}{3} \sum_{j=1}^{3} \pi_{t-j} \right)^2 / 0.477^2 \right\},$$

$$R^2 = 0.606,\; se = 0.476,\; obs = 396,\; LM(1) = 0.146,\; LM(1\text{-}12) = 0.189$$

where t-statistics in absolute values are given in parentheses below the parameter estimates, R^2 denotes the coefficient of determination, se is the standard error of the regression, obs is the number of observations, $LM(1)$ and $LM(1\text{-}12)$ are p-values for Lagrange multiplier test statistics for first-order, and up to twelfth-order serial correlations in the residuals, respectively.

Note that the estimate of the scaling parameter γ is expressed in terms of the transition variable $z_t = 3^{-1} \sum_{j=1}^{3} \pi_{t-j}$ divided by its sample standard deviation 0.477.

The model performs well in terms of the goodness of fit and statistically significant coefficient estimates. Furthermore, there is no evidence of remaining autocorrelations in residuals.

Based on the parameter estimates, we show the implied ERPT $\widehat{\phi}_{2,0} + \widehat{\phi}_{4,0} G(z_t; \widehat{\gamma})$ in Figure 4 against the transition variable $z_t = 3^{-1} \sum_{j=1}^{3} \pi_{t-j}$ (the circles denoting the actual data points). The plot suggests that the degree of ERPT becomes largest when the transition variable, namely the average lagged inflation rate, exceeds 2 percent in absolute value. Figure 5 shows the smoothed estimates of the time-varying ERPT, based on the 12-month moving averages, along with their two-standard error bands. The smoothed plot illustrates three distinct high ERPT episodes. The first high ERPT period corresponds to the second oil shock in the late 1970s. During the 1980s and 1990s, the ERPT is relatively stable except for the early 1990s when the producer price index is relatively volatile. During the decade beginning in 2000, the ERPT becomes high again due to the increased inflation.

4.3 Symmetric DLSTAR model

To select the delay parameter for the transition variable and lags for the regressors in a symmetric version of the DLSTAR model, we use a procedure similar to the one employed for the ESTAR model estimation. We select $d = 1$ and the estimation results are given as follows:

$$
\begin{aligned}
\pi_t \quad = \quad & \underset{(3.466)}{0.098} + \underset{(3.866)}{0.208}\pi_{t-1} + \underset{(4.278)}{0.159}\pi_{t-3} - \underset{(2.195)}{0.101}\pi_{t-5} + \underset{(12.519)}{0.349}\,\Delta(s_t + p_t^*) \\
& + \underset{(2.341)}{0.075}\Delta(s_{t-1} + p_{t-1}^*) - \underset{(2.441)}{0.070}\Delta(s_{t-2} + p_{t-2}^*) + \underset{(2.081)}{0.066}\Delta(s_{t-5} + p_{t-5}^*) \\
& + \Big[\underset{(1.150)}{0.242}\pi_{t-4} - \underset{(2.749)}{0.739}\pi_{t-5} + \underset{(5.269)}{1.230}\pi_{t-6} + \underset{(23.333)}{0.651}\,\Delta(s_t + p_t^*) \\
& - \underset{(5.568)}{0.438}\Delta(s_{t-1} + p_{t-1}^*) + \underset{(3.109)}{0.350}\Delta(s_{t-2} + p_{t-2}^*) - \underset{(2.695)}{0.534}\Delta(s_{t-4} + p_{t-4}^*) \\
& - \underset{(1.957)}{0.356}\Delta(s_{t-5} + p_{t-5}^*)\Big] G\left(z_t; \widehat{\gamma}, \widehat{c}\right) + \widehat{\varepsilon}_t,
\end{aligned}
$$

$$
\begin{aligned}
G\left(z_t;\widehat{\gamma},\widehat{c}\right) &= \left(1+\exp\left\{\underset{(2.924)}{-5.130}\left(\pi_{t-1}-\underset{(21.283)}{1.474}\right)/0.686\right\}\right)^{-1} \\
&\quad +\left(1+\exp\left\{\underset{(2.924)}{5.130}\left(\pi_{t-1}+\underset{(21.283)}{1.474}\right)/0.686\right\}\right)^{-1},
\end{aligned}
$$

$$R^2=0.654,\ se=0.448,\ obs=396,\ LM(1)=0.040,\ LM(1\text{-}12)=0.242.$$

Again, the estimate of the scaling parameter γ $(=\gamma_1=\gamma_2)$ is expressed in terms of a normalized transition variable. As shown in Figure 6, the shape of the implied ERPT $\widehat{\phi}_{2,0}+\widehat{\phi}_{4,0}G(z_t;\widehat{\gamma},\widehat{c})$ as a function of the transition variable $z_t=\pi_{t-1}$ somewhat resembles the shape of the transition function of TAR model predicted by the two-period contract case (Figure 1). A threshold-model-like shape of the transition function results in many data points near the lowest ERPT. Because of this feature, the time series plot of ERPT based on the DLSTAR model shown in Figure 7 shows more observations of low and stable ERPT around 0.35 compared to the case of the ESTAR model.

4.4 Asymmetric DLSTAR model

We now turn to the estimation of the asymmetric version of the DLSTAR model to incorporate the possibility of asymmetric adjustment. Minimizing the sum of the squared residuals yields the choice of $d=1$. The final specification of the model with parameter estimates is as follows:

$$
\begin{aligned}
\pi_t &= \underset{(3.349)}{0.095}+\underset{(4.183)}{0.270}\pi_{t-1}+\underset{(4.094)}{0.153}\pi_{t-3}-\underset{(2.326)}{0.105}\pi_{t-5}+\underset{(12.352)}{0.341}\Delta(s_t+p_t^*) \\
&\quad +\underset{(1.879)}{0.062}\Delta(s_{t-1}+p_{t-1}^*)-\underset{(2.747)}{0.078}\Delta(s_{t-2}+p_{t-2}^*)+\underset{(2.071)}{0.064}\Delta(s_{t-5}+p_{t-5}^*) \\
&\quad +\Big[-\underset{(1.722)}{0.198}\pi_{t-1}-\underset{(1.979)}{0.510}\pi_{t-5}+\underset{(4.666)}{1.001}\pi_{t-6}+\underset{(23.868)}{0.659}\Delta(s_t+p_t^*) \\
&\quad -\underset{(3.324)}{0.338}\Delta(s_{t-1}+p_{t-1}^*)+\underset{(3.352)}{0.417}\Delta(s_{t-2}+p_{t-2}^*)-\underset{(2.685)}{0.298}\Delta(s_{t-4}+p_{t-4}^*) \\
&\quad -\underset{(2.699)}{0.482}\Delta(s_{t-5}+p_{t-5}^*)\Big]G\left(z_t;\widehat{\gamma}_1,\widehat{\gamma}_2,\widehat{c}_1,\widehat{c}_2\right)+\widehat{\varepsilon}_t,
\end{aligned}
$$

$$
\begin{aligned}
G\left(z_t;\widehat{\gamma}_1,\widehat{\gamma}_2,\widehat{c}_1,\widehat{c}_2\right) \;&=\; \left(1+\exp\left\{-\underset{(1.129)}{5.762}\left(\pi_{t-1}-\underset{(14.924)}{1.591}\right)/0.686\right\}\right)^{-1} \\
&\quad+\left(1+\exp\left\{\underset{(1.124)}{55.253}\left(\pi_{t-1}+\underset{(156.218)}{1.293}\right)/0.686\right\}\right)^{-1},
\end{aligned}
$$

$$
R^2 = 0.663,\ se = 0.443,\ obs = 396,\ LM(1) = 0.073,\ LM(\text{1-12}) = 0.247.
$$

Again, the estimates of the scaling parameters γ_1 and γ_2 are expressed in terms of the normalized transition variable.

Figure 8 plots the implied ERPT $\widehat{\phi}_{2,0}+\widehat{\phi}_{4,0}G(z_t;\widehat{\gamma}_1,\widehat{\gamma}_2,\widehat{c}_1,\widehat{c}_2)$ against the transition variable $z_t=\pi_{t-1}$ allowing for the asymmetric adjustment. In terms of the shape of the transition function, the asymmetric DLSTAR specification result is similar to that of the symmetric DLSTAR specification. However, because the estimate of γ_2 is much larger than that of γ_1, the transition is much faster in the negative region. Figure 9 shows the smoothed plots of the ERPT implied by the asymmetric DLSTAR model estimates over the sample period. The behavior of estimated ERPT is very similar to the one implied by the symmetric DLSTAR model.

4.5 Specification test

Table 2 reports the results of the specification test to select an appropriate transition function among the ESTAR, the symmetric DLSTAR and the asymmetric DLSTAR models. In some cases, the null hypothesis of the ESTAR model against the asymmetric DLSTAR model cannot be rejected (see F_3 with d greater than 3). On the other hand, the evidence suggests rejecting the symmetric DLSTAR model in favor of the asymmetric DLSTAR (F_{13}) and the ESTAR specifications ($F_{1|3}$). While the evidence is somewhat mixed, the ESTAR and asymmetric DLSTAR specifications may be slightly better than the symmetric DLSTAR specification.

5 Conclusion

In this paper, we show that the STAR models, the parsimonious parametric nonlinear time series models, offer a very convenient framework in examining the relationship between the ERPT and inflation. First, a simple theoretical model of ERPT determination suggests that the dynamics of ERPT can be well approximated by a class of STAR models with lagged inflation as a transition variable. Second, we can employ various U-shaped transition functions in the estimation of the time-varying ERPT. When this procedure is applied to US import and domestic price data, we find the supporting evidence of nonlinearity in ERPT dynamics. Our empirical results imply that the period of low ERPT is likely to be associated with the low inflation.

According to our model, the degree of ERPT varies over time because the fraction of importing firms opting out from the contract is endogenously determined by importing firms' optimization behavior. In the model, however, all imports are treated as if they are invoiced in the producer's (exporter's) currency. An alternative approach in introducing a time-varying ERPT is to use a model in which exporting firms endogenously choose between producer currency pricing (PCP) and local currency pricing (LCP). For example, a recent study by Gopinath et al. (2010) extends the model of Engel (2006) and investigates the role of the invoice currency in determining the observed ERPT. Our analysis does not consider this channel partly because we do not have data on individual exporters' invoice currency. Incorporating the effect of currency choice in our estimation procedure seems to be a promising direction for further analysis.

Appendix: A Model of Importers

In this appendix we provide a full description of the theoretical model and derive its implications discussed in Section 2. There is a continuum of monopolistically competitive importing firms, each of which imports a differentiated intermediate good from abroad and sells it to a representative domestic final good producer. In each time period, a constant fraction $1/N$ of all importing firms and the final good producer write their pricing contracts of N periods long. An importing firm that writes the pricing contract at time $t-j$ (for $j = 0, 1, ..., N-1$) and imports a good $i \in [0, 1]$, at time t is facing a demand given by

$$C_t(i, t-j) = \left(\frac{P_t(i, t-j)}{P_t(t-j)}\right)^{-\theta} C_t(t-j)$$

where $\theta > 1$ is a constant elasticity of substitution. Here, $P_t(i, t-j)$ is the price of a good i imported by a firm with a contract beginning in period $t-j$. $P_t(t-j) = \left(\int_0^1 P_t(i, t-j)^{1-\theta} di\right)^{1/(1-\theta)}$ is the price index for the composite intermediate good sold by importing firms whose contracts begin in period $t-j$. $C_t(t-j)$ is the demand for the corresponding composite good. The elasticity of substitution among composite intermediate goods sold by each fraction $1/N$ of all importing firms is assumed to be one, and thus aggregate price index at time t (in log) is $p_t = N^{-1} \sum_{j=0}^{N-1} p_t(t-j)$ where $p_t(t-j) = \ln P_t(t-j)$.

All the differentiated intermediate goods are imported at the same foreign currency price, P_t^*, which is beyond the control of importers. The importer's profit, in terms of the domestic currency, at time t is given by

$$\Pi_t(i, t-j) = P_t(i, t-j) C_t(i, t-j) - (1+\tau) S_t P_t^* C_t(i, t-j)$$

where S_t is the nominal exchange rate, and τ is the iceberg transportation cost the importer must bear. The importer's desired price, which maximizes the profit under flexible price economy, is

$$\widehat{P}_t(i, t-j) = \frac{\theta}{\theta-1}(1+\tau) S_t P_t^*$$

where $\theta/(\theta-1)$, and $(1+\tau)S_t P_t^*$ represent the mark-up and marginal cost, respectively. By taking a log of the desired price, which is same across all the importing firms ($\widehat{P}_t = \widehat{P}_t(i, t-j)$), we have $\widehat{p}_t = s_t + p_t^* + \mu$ where $s_t = \ln S_t$ and

$\mu = \ln(\theta/(\theta-1)) + \ln(1+\tau)$. Both s_t and p_t^* are assumed to follow (possibly mutually correlated) random walk processes with a variance of the sum of each increment, $\Delta(s_t + p_t^*)$, given by σ^2.

In the initial period of the contract, importers set the price at $\widehat{p}_t$. For the rest of the contract period, they fully index their initial price $\widehat{p}_t$ to the aggregate inflation rate given by $\pi_t = p_t - p_{t-1}$. Note that prices are indexed to inflation of the initial period only, instead of following the period-by-period lagged inflation indexation rule as in Christiano et al. (2005). While the latter pricing scheme can be also introduced in our model, the former assumption greatly simplifies the analysis.

In reality, contracts written for fixed periods can, in special circumstances, be re-negotiated. By paying a fixed cost, firms can opt out of the contract and reset their price at the desired level. For example, in Devereux and Siu (2007), each firm observes its fix cost, which is assumed to be i.i.d. across firms, after setting its (two-period) contract price. Consequently, the pricing in the second period becomes state-dependent with all firms facing the same probability of opting out in the second period. We also let firms make their decision in a sequential manner by assuming that the aggregate inflation is not observed by individual firms at the time of the contract. However, instead of formally deriving the state-dependent pricing solution, we follow Ball et al. (1988), Romer (1990), and Devereux and Yetman (2002, 2010), among others, and re-formulate the firm's optimization behavior so that the probability of (not) changing its price to the desired price level is endogenously determined. Let $\kappa^{(t)}$ be the conditional probability that a firm will not opt out of the contract, provided that the firm is in the contract in the current period. After setting the new contract price $\widehat{p}_t$ at t, the firms observe the aggregate inflation π_t and choose $\kappa^{(t)}$ to maximize their profit. As in Walsh (2003), we can rewrite the intertemporal profit maximization condition using the expected squared deviation of the actual price from the desired price in each period.

(A) Two-period contract case

When $N = 2$, an optimal value of $\kappa^{(t)}$ is selected by minimizing the expected loss function given by

$$\begin{aligned} L_t &= E_t\left[\beta\kappa^{(t)}(\widehat{p}_t + \pi_t - \widehat{p}_{t+1})^2\right] + \beta(1-\kappa^{(t)})F \\ &= \beta F - \beta(F - \sigma^2 - \pi_t^2)\kappa^{(t)} \end{aligned}$$

where β is a discount factor and F is a fixed cost. Here we exclude the possibility of

$F < \sigma^2$, since the loss is always minimized by setting $\kappa^{(t)} = 0$ in such a case. When $F \geq \sigma^2$, the firm selects $\kappa^{(t)} = 1$ if $\pi_t^2 \leq F - \sigma^2$ and $\kappa^{(t)} = 0$ if $\pi_t^2 > F - \sigma^2$. Thus, for the given values of F and σ^2, $\kappa^{(t)}$ is simply a function of π_t. When we use the same argument, for any firms entering into contracts at time $t - j$, $\kappa^{(t-j)}$ is a function of π_{t-j} given by $\kappa(\pi_{t-j}) = \mathbf{1}\{|\pi_{t-j}| \leq \sqrt{F - \sigma^2}\}$. Using the definition of the aggregate price index, we have

$$\begin{aligned} p_t &= \frac{1}{2}\left(p_t(t) + p_t(t-1)\right) \\ &= (s_t + p_t^* + \mu) - \frac{\kappa(\pi_{t-1})}{2}\Delta(s_t + p_t^*) + \frac{\kappa(\pi_{t-1})}{2}\pi_{t-1} \end{aligned}$$

since the firms with new contracts set their price $p_t(t)$ at the desired price, $\widehat{p}_t = s_t + p_t^* + \mu$, and the firms with contracts made in the previous period set their price $p_t(t-1)$ at $(1-\kappa(\pi_{t-1}))\widehat{p}_t + \kappa(\pi_{t-1})(\widehat{p}_{t-1} + \pi_{t-1})$. The inflation dynamics are written as

$$\pi_t = \left(1 - \frac{\kappa(\pi_{t-1})}{2}\right)\Delta(s_t + p_t^*) + \frac{\kappa(\pi_{t-2})}{2}\Delta(s_{t-1} + p_{t-1}^*) + \frac{\kappa(\pi_{t-1})}{2}\pi_{t-1} - \frac{\kappa(\pi_{t-2})}{2}\pi_{t-2}.$$

We follow Devereux and Yetman (2010), among others, and consider the (short-run) ERPT in terms of the first derivative of π_t with respect to $\Delta(s_t + p_t^*)$, or

$$ERPT = 1 - \frac{\kappa(\pi_{t-1})}{2},$$

which depends on the lagged inflation, π_{t-1}. When $-\sqrt{F - \sigma^2} \leq \pi_{t-1} \leq \sqrt{F - \sigma^2}$, $\kappa(\pi_{t-1})$ takes a value of one and the ERPT becomes 0.5. On the other hand, when $|\pi_{t-1}| > \sqrt{F - \sigma^2}$, the model predicts a full ERPT.

(B) Three-period contract case

When $N = 3$, the loss function becomes a quadratic function of $\kappa^{(t)}$ given by

$$\begin{aligned} L_t &= E_t\left[\beta\kappa^{(t)}(\widehat{p}_t + \pi_t - \widehat{p}_{t+1})^2 + (\beta\kappa^{(t)})^2(\widehat{p}_t + 2\pi_t - \widehat{p}_{t+2})^2\right] \\ &\quad + \beta(1 - \kappa^{(t)})(1 + \beta)F + \beta^2\kappa^{(t)}(1 - \kappa^{(t)})F \\ &= \beta(1+\beta)F - \beta(F - \sigma^2 - \pi_t^2)\kappa^{(t)} - \beta^2(F - 2\sigma^2 - 4\pi_t^2)(\kappa^{(t)})^2. \end{aligned}$$

The first order condition yields the optimal $\kappa^{(t)}$ given by

$$\kappa(\pi_t) = \frac{-(F - \sigma^2 - \pi_t^2)}{2\beta(F - 2\sigma^2 - 4\pi_t^2)}$$

provided $F - \sigma^2 - \pi_t^2 > 0$ and $(F - \sigma^2 - \pi_t^2) + 2\beta(F - 2\sigma^2 - 4\pi_t^2) < 0$. In this case, $\kappa^{(t)}$ is a smooth function of the inflation rate π_t. Otherwise, $\kappa^{(t)}$ becomes a corner solution taking a value of either 0 or 1. In particular, if $F-\sigma^2-\pi_t^2>0$ and $(F-\sigma^2-\pi_t^2)+2\beta(F-2\sigma^2-4\pi_t^2)\geq 0$, then $\kappa(\pi_t)=1$. If $F-\sigma^2-\pi_t^2\leq 0$, then $\kappa(\pi_t)=0$. The aggregate price is given by

$$\begin{aligned}
p_t &= \frac{1}{3}\left(p_t(t)+p_t(t-1)+p_t(t-2)\right) \\
&= (s_t+p_t^*)-\frac{\kappa(\pi_{t-1})+\kappa(\pi_{t-2})^2}{3}\Delta(s_t+p_t^*)-\frac{\kappa(\pi_{t-2})^2}{3}\Delta(s_{t-1}+p_{t-1}^*) \\
&\quad +\frac{\kappa(\pi_{t-1})}{3}\pi_{t-1}+\frac{2\kappa(\pi_{t-2})^2}{3}\pi_{t-2}
\end{aligned}$$

where the second equality follows from $p_t(t-1)=(1-\kappa(\pi_{t-1}))\widehat{p}_t+\kappa(\pi_{t-1})(\widehat{p}_{t-1}+\pi_{t-1})$ and $p_t(t-2)=(1-\kappa(\pi_{t-2})^2)\widehat{p}_t+\kappa(\pi_{t-2})^2(\widehat{p}_{t-2}+2\pi_{t-2})$. The inflation dynamics are given by

$$\begin{aligned}
\pi_t &= \left(1-\frac{\kappa(\pi_{t-1})+\kappa(\pi_{t-2})^2}{3}\right)\Delta(s_t+p_t^*) \\
&\quad -\frac{1}{3}\left(\kappa(\pi_{t-2})^2-\kappa(\pi_{t-2})-\kappa(\pi_{t-3})^2\right)\Delta(s_{t-1}+p_{t-1}^*)+\frac{\kappa(\pi_{t-3})^2}{3}\Delta(s_{t-2}+p_{t-2}^*) \\
&\quad +\frac{\kappa(\pi_{t-1})}{3}\pi_{t-1}+\frac{1}{3}\left(2\kappa(\pi_{t-2})^2-\kappa(\pi_{t-2})\right)\pi_{t-2}-\frac{2\kappa(\pi_{t-3})^2}{3}\pi_{t-3}.
\end{aligned}$$

The ERPT is given by

$$ERPT=1-\frac{\kappa(\pi_{t-1})+\kappa(\pi_{t-2})^2}{3}$$

which now depends on π_{t-1} and π_{t-2}.

(C) N-period contract case

With a similar argument, for general N, the current inflation becomes a function of π_{t-j} for $j=1,...,N$ and $\Delta(s_{t-j}+p_{t-j}^*)$ for $j=0,...,N-1$. The ERPT for any N is given by

$$ERPT=1-\frac{\sum_{j=1}^{N-1}\kappa(\pi_{t-j})^j}{N}$$

where $\kappa(\pi_{t-j})$ is a nonlinear function of π_{t-j}. The second term $N^{-1}\sum_{j=1}^{N-1}\kappa(\pi_{t-j})^j$ represents the fraction of firms adapting the indexation rule and the ERPT can now vary from $1/N$ to 1. In general, the ERPT is a smooth nonlinear function of lagged inflation rates, with its dynamics possibly approximated by STAR models with a U-shaped transition function.

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Table 1. Tests for linearity against STAR models

		Transition Variable ($z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$)					
Test Statistics	H_0	$d = 1$	$d = 2$	$d = 3$	$d = 4$	$d = 5$	$d = 6$
LM	Linear AR	137.09	116.63	106.81	92.74	86.06	67.62
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
F_L	Linear AR	3.33	2.83	2.59	2.25	2.09	1.64
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)
LM^*	Linear AR	351.8	355.3	354.1	357.7	358.2	358.4
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

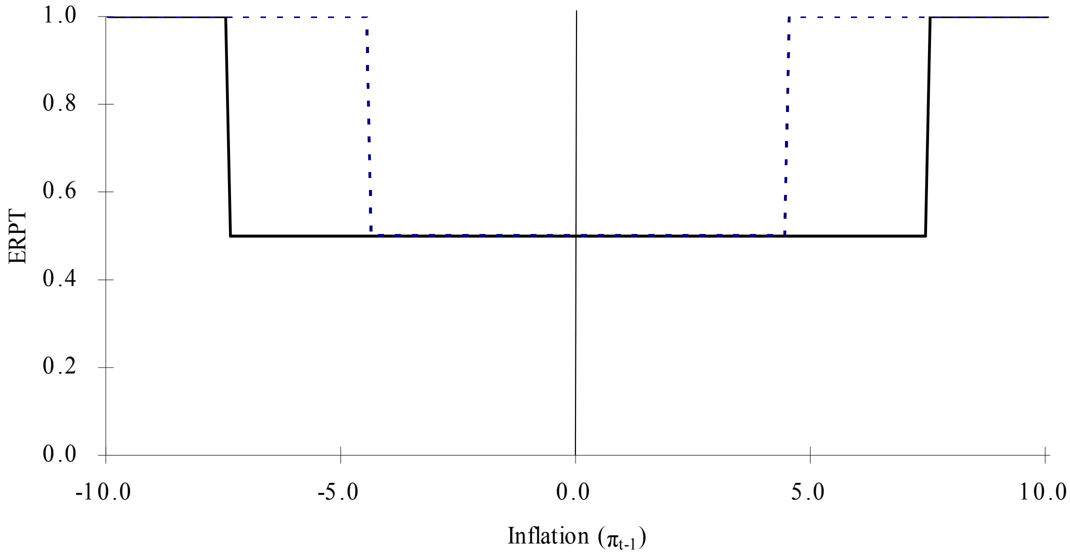
Notes: Lag length is $N = 6$. The LM test statistics, F version of the LM test statistics and the heteroskedasticity-robust variants of the LM test statistics are denoted as LM, F_L and LM^*, respectively. The numbers in parentheses below statistics are p-values.

Table 2. Specification tests for STAR models

		Transition Variable ($z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$)					
Test Statistics	H_0 (H_1)	$d = 1$	$d = 2$	$d = 3$	$d = 4$	$d = 5$	$d = 6$
F_3	ESTAR	3.82	2.91	2.31	1.49	1.70	1.14
	(Asymmetric DLSTAR)	(0.00)	(0.00)	(0.01)	(0.13)	(0.06)	(0.32)
$F_{1\|3}$	Symmetric DLSTAR	2.70	3.00	2.74	2.97	2.28	1.27
	(ESTAR)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.23)
F_{13}	Symmetric DLSTAR	3.04	2.76	2.37	2.10	1.89	1.16
	(Asymmetric DLSTAR)	(0.00)	(0.00)	(0.00)	(0.07)	(0.02)	(0.29)

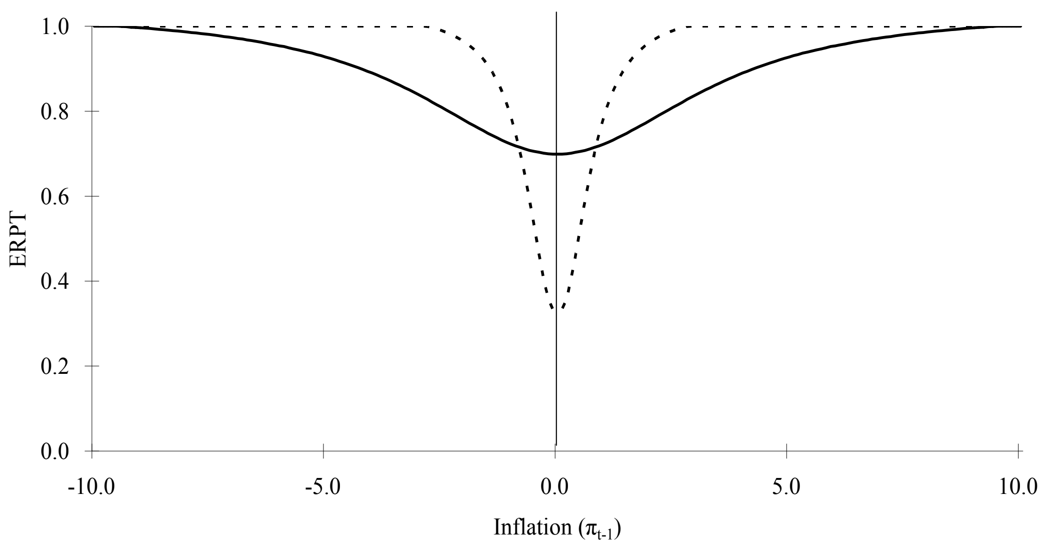
Notes: Lag length is $N = 6$. F_3 is the F test statistic for $H_0 : \beta_3 = 0$ against $H_1 : \beta_3 \neq 0$. $F_{1|3}$ is the F test statistic for $H_0 : \beta_1 = 0|\beta_3 = 0$ against $H_1 : \beta_1 \neq 0|\beta_3 = 0$. F_{13} is the F test statistic for $H_0 : \beta_1 = \beta_3 = 0$ against $H_1 : \beta_1 \neq 0$ and $\beta_3 \neq 0$. Under the null hypothesis, three sets of F test statistics follow F distributions with $(2N, T - 8N - 1)$, $(2N, T - 6N - 1)$ and $(4N, T - 8N - 1)$ degrees of freedom, respectively. The numbers in parentheses below F statistics are p-values.

Figure 1. ERPT and inflation: Two-period contract case (N=2)



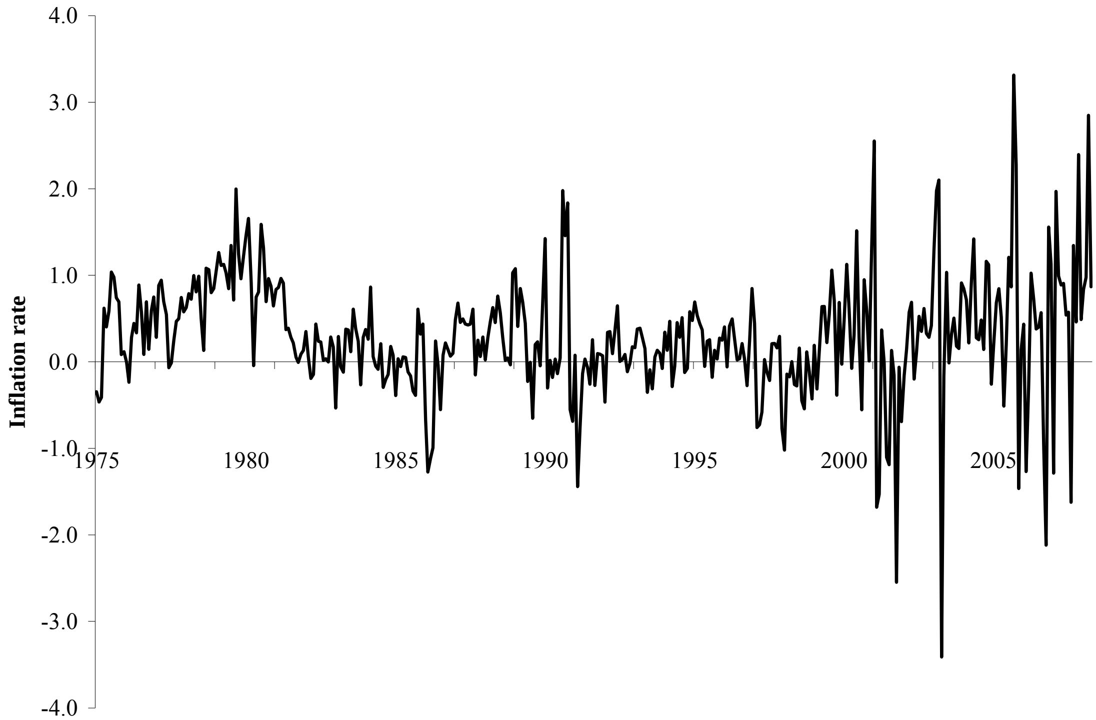
Notes: Solid line: F = 155 and σ^2 = 100. Dotted line: F = 120 and σ^2 = 100.

Figure 2. ERPT and inflation: Three-period contract case (N=3)



Notes: Solid line: F = 260 and σ^2 = 170. Dotted line: F = 20 and σ^2 = 12.

Figure 3. Producer price index inflation



Note: Seasonally adjusted series.

Figure 4. ERPT against the transition variable: ESTAR model

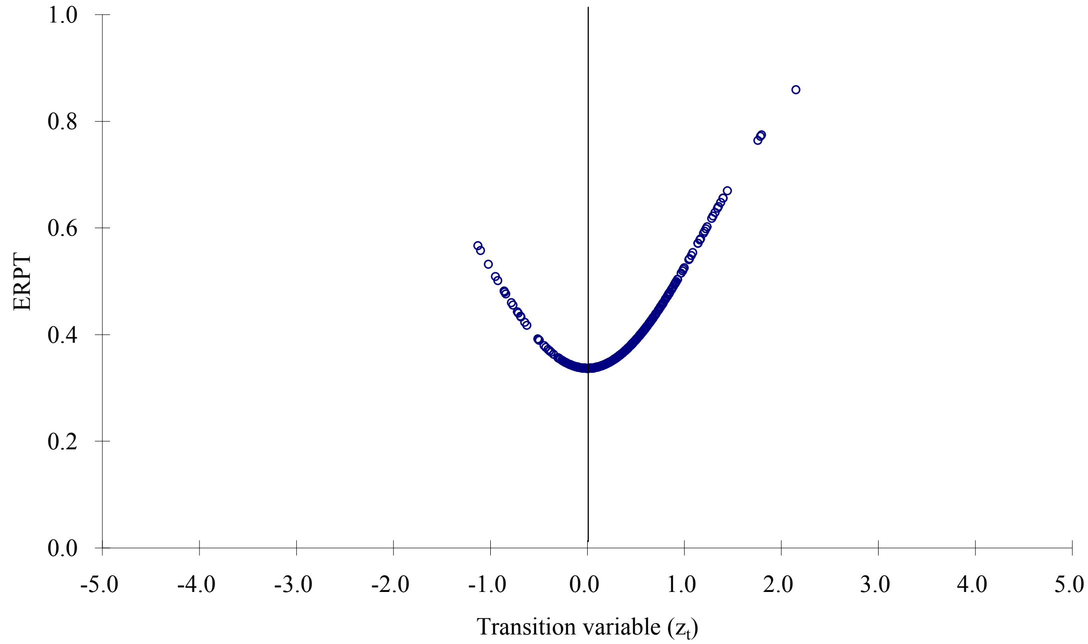


Figure 5. ERPT over time: ESTAR model

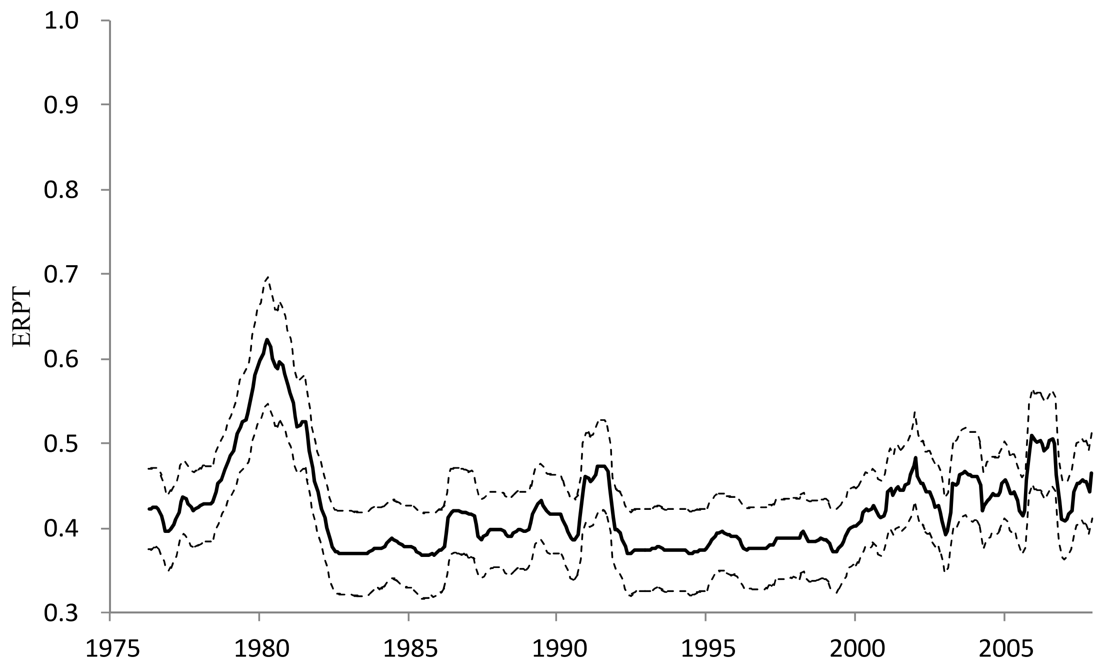


Figure 6. ERPT against the transition variable: Symmetric DLSTAR model

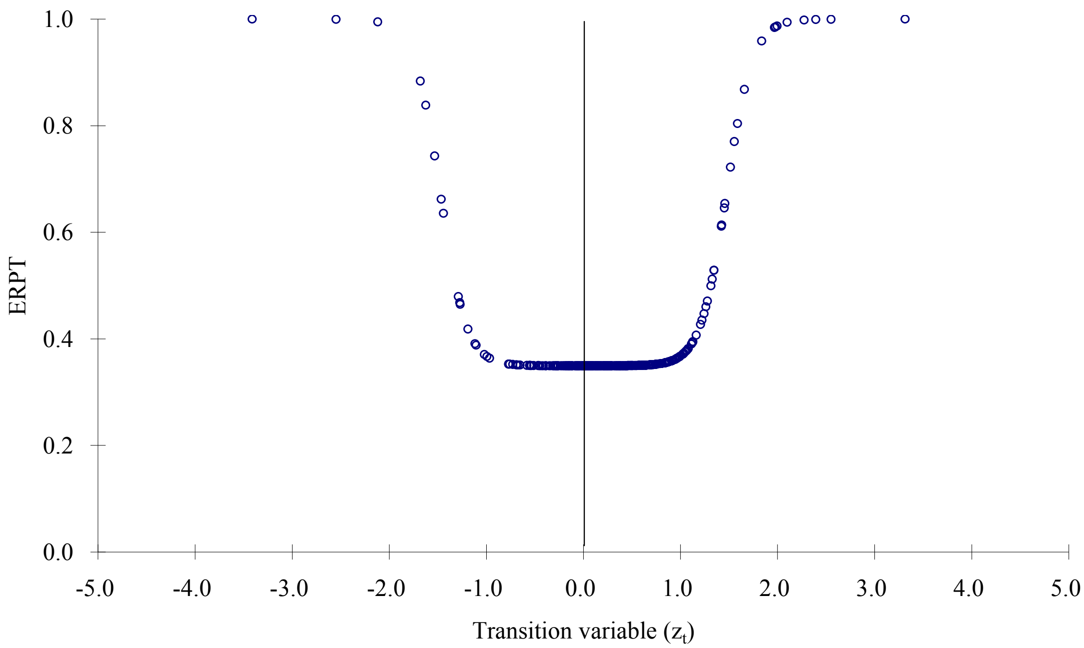


Figure 7. ERPT over time: Symmetric DLSTAR model

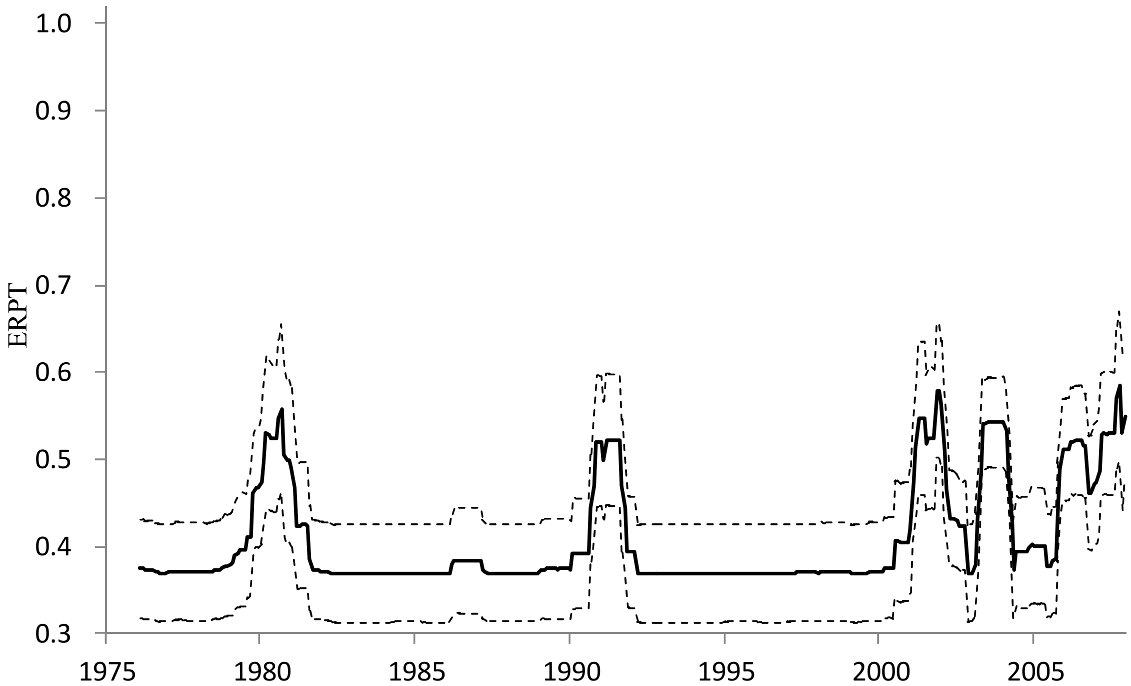


Figure 8. ERPT against the transition variable: Asymmetric DLSTAR model

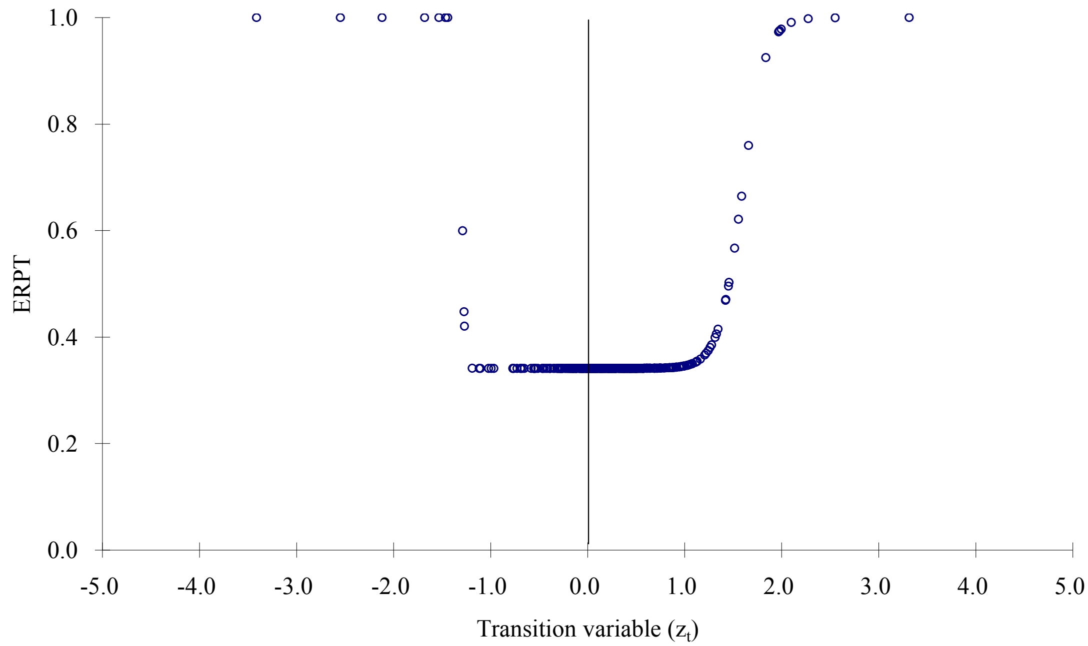


Figure 9. ERPT over time: Asymmetric DLSTAR model