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. Using 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 with inflation being a transition variable. We employ several U-shaped transition functions in the estimation of the time-varying ERPT to U.S. domestic prices. The result suggests that declines in the ERPT during the 1980s and the 1990s are associated with the 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 that determine the size of spill-over effects of monetary policy. Over the past decade, a number of empirical studies investigated whether ERPT decreased during the 1980s and the 1990s.[1] If the reduction of ERPT is true, it is natural to conjecture the lower and more stable inflations as a possible factor because the timing corresponds to the period of price stability in many countries. 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 to domestic prices and inflation using nonlinear time series methods. In particular, we employ the class of smooth transition autoregressive (STAR) models to describe the case when the degree of ERPT is determined by the level of the lagged inflation rate. Most of previous empirical studies on the positive association between the ERPT and inflation have been focusing on the cross-country evidence, including the analyses by Calvo and Reinhart (2002), Choudhri and Hakura (2006) and Devereux and Yetman (2008). Instead of examining the relationship between the ERPT and average inflation rate across countries, we are interested in examining the role of inflation on the time-varying ERPT using a time series modeling framework.

In the empirical literature on the nonlinear adjustment of real exchange rates, the STAR models have been popularly employed by many studies, including Michael, Nobay and Peel (1997), Taylor and Peel (2000), Taylor, Peel and Sarno (2001), and Kilian and Taylor (2003), among others. However, this time series approach has rarely

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

been used in the studies of ERPT.[2] We employ several U-shaped transition functions in the STAR model to consider alternative forms of adjustment. Our method is applied to monthly U.S. import and domestic price data to evaluate the time-varying structure of ERPT during the period from 1975 to 2007.

To motivate our nonlinear regression approach, we first construct a simple theoretical model of importing firms where ERPT is predicted to be a nonlinear function of the past inflation rate. The structure of our model is closely related to a model of ERPT developed by Devereux and Yetman (2008). As in Ball, Mankiw and Romer (1988), price setters are allowed to endogenously select the probability of adjusting their price to an optimal level. However, our model differs from that of Devereux and Yetman (2008) in several aspects. First, instead of all firms facing a infinite horizon profit maximization problem à la Calvo (1983), a fraction of firms make finite-period contracts every period as in the Taylor (1980) type staggered contract. Second, each pricing cohort of firms faces the problem of selecting a probability of opting out from a contract inflation indexation rule and setting an optimal price with a payment of a fixed cost. With these modifications, our model predicts that ERPT depends on the lagged inflation unlike the case of Devereux and Yetman (2008) where ERPT depends on the steady-state inflation level of the economy. Using this model, we show that dynamics of ERPT predicted by the model can be well-approximated by using the STAR formulation. Our estimates of STAR models suggest that the past decline during the 1980s and the 1990s and recent increase in the ERPT to U.S. domestic price are well-explained by the changes in inflation.

The remainder of the paper is organized as follows. Section 2 describes the theoretical model and empirical model. Section 3 explains the data and provides the estimation results. Some concluding remarks are made in Section 4.

[2] One of the few exceptions is Herzberg, Kapetanios and Price's (2003) study of UK import price. However, they did not find the supporting evidence on nonlinearity.

2 Exchange Rate Pass-through into Domestic Prices

2.1 A Simple Model of Importers

As in the model of Devereux and Yetman (2008), we consider a continuum of importing firms, each of which purchases a differentiated good from abroad and sells it in the domestic market as a monopolistic competitor. To describe the price setting behavior of importing firms, we modify the Taylor's (1980) staggered pricing model in which contracts between the importers and the retailers are $N(\geq 2)$ periods long and a constant fraction $1/N$ of all importing firms determine their contracts in any given time period. A firm that determines the pricing contract at time $t-j$ (for $j = 0, 1, ..., N-1$) and imports a good i 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, $P_t(i, t-j)$ 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 aggregate price index for goods sold by all the importing firms with contracts beginning in period $t-j$ and $C_t(t-j)$ is the aggregate demand for the same importers. The prices of all differentiated imported goods, $i \in [0, 1]$, have the same foreign currency price, P_t^*. The firm sets prices in term of the domestic currency with its profit 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 importing firms must bear. The firm'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 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 $\mu = \ln(\theta/(\theta-1)) + \ln(1+\tau)$.[3] In this paper, 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. We further assume that the elasticity of substitution among aggregate consumption goods sold by each fraction $1/N$ of all importing firms is one, and thus aggregate price index at time t is given by $P_t = \left(\Pi_{j=0}^{N-1} P_t(t-j)\right)^{1/N}$, or $p_t = N^{-1}\sum_{j=0}^{N-1} p_t(t-j)$ where $p_t(t-j) = \ln P_t(t-j)$. Our interest is the effect of marginal cost changes $\Delta(s_t + p_t^*)$ on the aggregate inflation rate given by $\pi_t = p_t - p_{t-1}$.[4]

In reality, contracts written for fixed periods can be re-negotiated in special circumstances. Ball and Mankiw (1994) and Devereux and Siu (2007) add to the two-period Taylor model the possibility that in the second half of their contracts, firms can choose to opt out and reset the price by paying a fixed (menu) cost. We introduce a similar pricing scheme by dividing the total of N periods into two subperiods with the first subperiod $N_1(\geq 1)$ during which the firms follow the contract pricing rule, and the second subperiod $N_2(= N - N_1)$ after opting out from the contract. During the contract period, firms are assumed to fully index their prices to aggregate inflation of the initial period. 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.[5] 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. Since the marginal cost follows a random walk and the probability of opting out is not

[3] In our empirical part of the analysis, we examine not only the case with constant μ but also the case where μ varies in response to demand shocks.

[4] As in Devereux and Yetman (2008), a full pass-through of nominal exchange rate changes into the import prices is implicitly assumed here. Recent empirical study by Goldberg and Campa (2008) also investigates the pass-through into the domestic prices rather than import prices using a model of the production sector with imported inputs.

[5] For this reason, Devereux and Siu (2007) refer it to the hybrid time- and state-dependent pricing rule.

known in the beginning, all the firms entering into new contracts at time t set their price at $\widehat{p}_t$. The firm that decides to opt out at time $t+N_1$ also chooses a flexible price at each period during the second subperiod with its entire price path given by $\{\widehat{p}_t, \widehat{p}_t+\pi_t, ..., \widehat{p}_t+(N_1-1)\pi_t, \widehat{p}_{t+N_1}, ..., \widehat{p}_{t+(N-1)}\}$.[6]

In this paper, we do not formally derive the state-dependent pricing solution. Instead, we follow Ball, Mankiw and Romer (1988), Romer (1990) and Devereux and Yetman (2002, 2008), 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 probability that a firm under the contract in the current period to maintain the contract price in the next period.[7] Here, a superscript t in parenthesis signifies that this probability applies to all the firms entering into new contracts at time t, but not to other pricing cohorts. After setting the new contract price 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. In our case, an optimal value of $\kappa^{(t)}$ is selected by minimizing the expected loss function

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

where β is a discount factor and F is a fixed cost. The above function implies that the loss is an increasing function of the squared inflation rate, π_t^2. As the inflation rate rises (relative to the size of the fixed cost), the firm can minimize the loss by avoiding the inflation indexation. This leads to a lower $\kappa^{(t)}$ or a shorter average length of N_1

[6] Note that prices are indexed to inflation of the initial period only instead of using the period-by-period lagged inflation indexation rule such as the one by Christiano, Eichenbaum and Evans (2005). While the latter pricing scheme can be also introduced in our model, the former assumption greatly simplifies the analysis.

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given by $\sum_{\ell=1}^{N-1}(\kappa^{(t)})^{\ell}+1$. In an extreme case of a high inflation, $\kappa^{(t)}=0$ (or $N_1=1$) is selected with a pricing path given by $\{\widehat{p}_t, \widehat{p}_{t+1}, ..., \widehat{p}_{t+(N-1)}\}$. In the other extreme case of a large fixed cost, $\kappa^{(t)}=1$ (or $N_1=N$) is selected with a pricing path given by $\{\widehat{p}_t, \widehat{p}_t+\pi_t, \widehat{p}_t+2\pi_t, ..., \widehat{p}_t+(N-1)\pi_t\}$.

While the formulation of our model closely follows that of Devereux and Yetman (2008), there are at least two notable differences between the pricing schemes of two models. First, importers in Devereux and Yetman (2008) face the problem of choosing the probability κ of making no price adjustment to maximize the profit over the infinite horizon under the framework of a standard Calvo (1983) type sticky price model. In contrast, our importers optimize over the finite contract period in the Taylor (1980) type staggered pricing model by choosing the probability $\kappa^{(t)}$ of maintaining the contract pricing rule. Since the optimization problem is resolved by importers entering into new contracts every period, the selected probability $\kappa^{(t)}$ generally differs across pricing cohorts, unlike κ which is common for all the firms in the economy.[8] Second, our importers are assumed to adapt an inflation indexation rule during the contract period, rather than to fix their price at a constant level. Under such a contract, price deviations from the desired price level depend on the aggregate inflation rate of the initial contract period. Thus, the optimal level of the probability $\kappa^{(t)}$ becomes a function of inflation. Because of this mechanism, our model predicts that ERPT depends on the lagged inflation unlike the case of Devereux and Yetman (2008) where ERPT depends on the steady-state inflation level of the economy. Below we derive the functional form of the ERPT, and show its dependence on the multiple pricing cohorts in the economy.

(A) Two-period contract case

Let us first consider the simplest case where each contract is written for two

[8] It should also be noted that $1-\kappa$ in the Calvo type model is an *unconditional* probability of the firm setting the optimal price in each period, but $1-\kappa^{(t)}$ in our model is a *conditional* probability of adapting the optimal pricing rule given that the firm has been following the contract pricing rule.

periods as in Ball and Mankiw (1994) and Devereux and Siu (2007). The loss function (1) with $N = 2$ is given by

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

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 given values of F and σ^2, $\kappa^{(t)}$ is simply a function of π_t. Using 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

$$\kappa(\pi_{t-j}) = \begin{cases} 1 & \text{if} \quad -\sqrt{F-\sigma^2} \leq \pi_{t-j} \leq \sqrt{F-\sigma^2} \\ 0 & \qquad\qquad\text{otherwise} \end{cases}$$

From the definition of the aggregate price index, we have

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

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 is written as

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

We follow Devereux and Yetman (2008) 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}.[9] 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

[9] Because of the random walk assumption in our model, this ERPT also corresponds to the ERPT defined by the first derivative of π_t with respect to Δs_t.

$|\pi_{t-1}| > \sqrt{F-\sigma^2}$, the model predicts a full ERPT. In summary, the model with $N = 2$ implies the abrupt transition of the degree of the ERPT depending on the relative size of a threshold variable $|\pi_{t-1}|$ and a threshold value $\sqrt{F-\sigma^2}$ as depicted in Figure 1. The shape of the step function in the figure suggests the possibility of approximating the ERPT by a variation of a threshold autoregressive model (TAR), sometimes referred to as the three-regime TAR model or the band TAR model.

(B) Three-period contract case

When $N = 3$, the loss function (1) 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.[10] 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

[10] 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 $\kappa(\pi_t) = 0$.

are given by

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

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}. The ERPT takes a minimum value of 1/3 when both $\kappa(\pi_{t-1})$ and $\kappa(\pi_{t-2})$ are one. In the other extreme case, a full ERPT can be obtained when both $\kappa(\pi_{t-1})$ and $\kappa(\pi_{t-2})$ are zero. Figure 2 shows the relationship between the inflation rate (imposing $\pi_{t-1} = \pi_{t-2}$) and the ERPT when the ERPT takes the value between zero and one. The smooth nonlinear relationship between the inflation and ERPT resembles the adjustment dynamics described by the class of STAR model with the lagged inflation rates used as transition variables. In particular, the symmetric relationship around zero suggests a symmetric U-shaped transition function such as an exponential function.

(C) N-period contract case

A similar argument yields the ERPT for any N 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. Again, the ERPT is a smooth nonlinear function of inflation, with its dynamics possibly approximated by a U-shaped transition function with a set of lagged inflation rates used as transition variables. The current inflation becomes a function of π_{t-j} for $j = 1, ..., N$ and $\Delta(s_{t-j} + p_{t-j}^*)$ for $j = 0, ..., N-1$.

2.2 STAR Models

To seek for a suitable specification of the STAR model used in the empirical analysis, let us summarize the predictions of the theoretical model introduced in the previous subsection. First, higher inflation (in absolute value) results in higher degree of the ERPT. Second, the ERPT is a symmetric function of the past inflation rates around zero. Finally, in general, dynamics of the ERPT can be described by a smooth rather than an abrupt transition using past inflation rates as transition variables possibly with multiple lags. Only exception is a special case of two-period contract which predicts a discrete transition typically assumed in the TAR model.

To capture these features in a parsimonious parametric model, we primarily employ the exponential STAR (ESTAR) model with an exponential function used as a symmetric U-shaped transition function. 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). Since our objective is to find the relationship between π_t and $\Delta(s_t + p_t^*)$, we estimate a bivariate version of the ESTAR model specified as

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

where $\varepsilon_t \sim$i.i.d.$(0, \sigma_\varepsilon^2)$ and an exponential transition function given by

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

where $\gamma(> 0)$ is a parameter defining the smoothness of the transition, z_t is a transition variable. While our theoretical model suggests a combination of past inflation rates as transition variables, here we consider a parsimonious specification and use a moving average of the past inflation rates as a single transition variable,

$z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$.[11] 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 STAR model based on a different U-shaped transition function constructed from the combination of two logistic functions. This variant of a logistic STAR (LSTAR) model has been considered in Granger and Teräsvirta (1993) and Bec, Ben Salem, and Carrasco (2004) and sometimes referred to as the three regime LSTAR model. Here we simply call the model as a dual (or double) LSTAR (DLSTAR) model to emphasize the presence of two logistic functions.[12] The transition function in the DLSTAR model is given by

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

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 definition 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, Teräsvirta, and Franses (2002), the transition function in the ESTAR model collapses to a constant when γ approaches

[11] 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 one.

[12] We use this terminology since the model differs from the multple regime STAR models defined in van Dijk, Teräsvirta, and Franses (2002).

infinity. Thus the model does not nest the TAR model with abrupt transition predicted by the theory when there are only two generations of firms in the economy. In contrast, DLSTAR model includes the TAR model by letting γ_1, γ_2 tend to infinity. Second, and more importantly, the model can incorporate both symmetric ($\gamma_1 = \gamma_2$ and $c_1 = c_2$) 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 which predicts a symmetric relationship between the ERPT and the inflation rate. In the following section, we mainly employ the symmetric DLSTAR by setting $\gamma_1 = \gamma_2 = \gamma$ and $c_1 = c = c$ but also consider the possibility of the asymmetric DLSTAR models.

3 Empirical Results

3.1 Main Results

All the data we use in the STAR estimation is the taken from *International Financial Statistics* (IFS) of 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 U.S. import price index constructed by Bureau of Labor Statistics is based on U.S. dollar prices paid by the U.S. importer, it is simply computed as $\Delta(s_t + p_t^*) = 1200 \times (\ln IMP_t - \ln IMP_{t-1})$ where IMP_t is the import price after making a seasonal adjustment using X-12-ARIMA procedure. The prices are generally either "free on board (f.o.b.)" foreign port or "cost, insurance, and freight (c.i.f.)" U.S. port transaction prices, depending on the practices of the individual industry. In either case, under our assumption of constant iceberg transaction cost (proportional to import price in domestic currency), the same formula can be used to compute the monthly log changes in 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 consumer price index since domestic price in our model is the price at which the importers sell their imports to domestic distributors rather than the retail price.[13] The monthly log inflation is computed as $\pi_t = 1200 \times (\ln PPI_t - \ln PPI_{t-1})$ where PPI_t is the seasonally adjusted U.S. 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.

Before estimating the ESTAR model, we first conduct the LM tests of linearity against ESTAR alternative developed by Saikkonen and Luukkonen (1988). See Appendix A.1 for the details of this linearity test. We will denote the original LM test by LM_1 and its heteroskedasticity-robust variant suggested by Granger and Teräsvirta (1993) by LM_1^*. The results using $N = 13$ and d between 1 and 6 are reported in the upper half of Table 1. Both tests strongly suggest the presence of nonlinearity in inflation dynamics for all values of d.

Second, we search for the length of moving average d in the transition variable z_t that best fits the specification. We fix the lag length $N = 13$ and search for the value of d between 1 and 12 which minimizes the residual sum of squares from the nonlinear least squares regression of (4). This leads to the choice of $d = 3$.

Third, we adopt a general-to-specific approach suggested by van Dijk, Teräsvirta, and Franses (2002), in arriving at the final specification. Starting with a model with $N = 13$, we sequentially remove the lagged variables for which the t statistic of the corresponding parameter is less than 1.5 in absolute value. The resulting final specification and the estimates for the ESTAR model are provided in the panel A of Table 2. The estimate of the scaling parameter γ is expressed in the form divided by the sample standard deviation of the transition variable z_t which is 0.477. The model performs well in terms of the goodness of fits and statistically significant coefficient estimates. Furthermore, there is no evidence of remaining autocorrelations

[13] There are other studies that also reports ERPT to producer price index. See, for example, Choudhri, Faruqee and Hakura (2005) and McCarthy (2007).

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 the degree of ERPT becomes the largest when the transition variable, namely the average inflation rate, becomes above 20% in absolute term. Figure 5 shows the time-varying ERPT implied by the estimates over the sample period. The ERPT takes the values between 0.34 and 0.85. The plots illustrate three distinct high ERPT episodes. The first high ERPT period corresponds to the second oil shock in the 1970s. During the 1980s and the 1990s, the ERPT is relatively stable except for the early 1990s when the producer price index is relatively volatile. During the 2000s, the ERPT becomes high again due to the increased volatility of inflation.

The panel B of Table 2 shows the estimation result of the symmetric DLSTAR model. Using the procedure similar to the one employed for the ESTAR model estimation, $d = 1$ for the transition variable and lags for the regressors are selected. Again, the estimate of the scaling parameter γ $(= \gamma_1 = \gamma_2)$ is expressed in the form divided by the sample standard deviation of the 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). In addition to a threshold model-like shape of the transition function, a larger variation of the transition variable, due to one lagged inflation $(d = 1)$ instead of its smoothed average $(d > 1)$, results in many data points near full ERPT. Because of these features, the time series variation of ERPT based on the DLSTAR model shown in Figure 7 becomes larger than the one based on the ESTAR model shown in Figure 5.

3.2 Further Analyses

(A) Introduction of asymmetric adjustment

We now turn to the estimation of the asymmetric DLSTAR model to incorporate the possibility of asymmetric adjustment. Let us here consider the linearity test which has a power against asymmetric DLSTAR alternative. Since such a test is not available in the literature, we provide the details of the construction of this linearity test in Appendix A.1. The lower half of Table 1 provides the evidence against linearity in favor of the asymmetric DLSTAR specification using the standard test LM_2 and its heteroskedasticity-robust variant LM_2^*. Minimizing the sum of the squared residuals yields the choice of $d = 1$. The final specification of the model with parameter estimates are presented in Table 3. Again, the estimates of the scaling parameters γ_1 and γ_2 are expressed in the form divided by the sample standard deviation of the 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 $z_t = \pi_{t-1}$. In terms of the shape of the transition function, the asymmetric DLSTAR specification yields a very similar result compared to the symmetric DLSTAR specification. However, because the estimate of γ_1 is much higher than that of γ_2, the transition is much faster in the negative region. Figure 9 shows the time series plots of the ERPT implied by the asymmetric DLSTAR model estimates over the sample period. The ERPT varies between the values of 0.34 and 1.00 and its path is very similar to the one implied by the symmetric DLSTAR model.

(B) Specification test for the choice of the transition function

Table 4 reports the results of the specification test to select an appropriate transition function among the ESTAR, the symmetric DLSTAR and the asymmetric DLSTAR models. The details of the test statistics are also provided in Appendix A.2. In some cases, the null hypothesis of the ESTAR model against the asymmetric DLSTAR model cannot be rejected (see LM_4 with d greater than 3). On the other hand, the evidence strongly suggests the rejection of the symmetric DLSTAR model in favor of the asymmetric DLSTAR (LM_3 and LM_3^*) and the ESTAR specifications (LM_5 and LM_5^*). Overall, the results weakly suggest either the ESTAR or the asymmetric DLSTAR model over the symmetric DLSTAR model.

(C) Introduction of demand shocks

Following Campa and Goldberg (2005), we also include the real output measure as an additional variable which represents the demand shock component of the domestic market. We use the index of industrial production normalized by the consumer price index (CPI) as a proxy for the change in demand. A modified ESTAR model we estimate is

$$\pi_t = F(\Delta(s_t+p_t^*), ..., \Delta(s_{t-N+1}+p_{t-N+1}^*), \pi_{t-1}, ..., \pi_{t-N}, z_t; \phi, \gamma) + \sum_{j=0}^{p} a_j \Delta iip_{t-j} + \varepsilon_t \tag{5}$$

where $F(\Delta(s_t + p_t^*), ..., \Delta(s_{t-N+1} + p_{t-N+1}^*), \pi_{t-1}, ..., \pi_{t-N}, z_t; \phi, \gamma)$ is the ESTAR function part used to obtain the main result, and Δiip_t is the first difference of logs of real industrial production multiplied by 1200. The lag length p is selected using the same method we use for selecting lags in the ESTAR part. The result is presented in Table 5. The time series path of ERPT is plotted in Figure 10. Overall, the performance of this extended ESTAR model is as good as that of the benchmark ESTAR model.

4 Conclusion

In this paper, we show that the STAR models, a parsimonious parametric nonlinear time series model, 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 inflation being 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 U.S. import and domestic price data, we find the supporting evidence of nonlinearities in inflation 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 to introduce a time-varying ERPT is to used a model in which exporting firms endogenously choose between the producer currency pricing (PCP) and local currency pricing (LCP). For example, a recent study by Gopinath, Itskhoki and Rigobon (2007) extends the model of Engel (2006) and investigates the role of the invoice currency in determining the observed ERPT. Our analysis do 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.1. Linearity test against DLSTAR model

Saikkonen and Luukkonen (1988) and Teräsvirta (1994) developed a methodology to test whether a series exhibits nonlinear behaviors described by STAR models. The test is based on a Taylor series approximation of the transition function of the STAR models.

Consider the following STAR model

$$y_t = x_t^{'}\phi_1 + G(z_t;\gamma)x_t^{'}\phi_2$$

where x_t is a vector of explanatory variables. Saikkonen and Luukkonen (1988) shows that for the ESTAR model, he replaces $G(z_t;\gamma) = 1 - \exp\{-\gamma z_t^2\}$ by a third-order Taylor series approximation with respect to γ evaluated at zero. This yields the following auxiliary regression:

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t\beta_2 + x_t^{'}z_t^2\beta_3 + e_t$$

Therefore, the linearity test against the ESTAR model is same as testing the joint restriction that all nonlinear terms are zero: $\beta_2 = \beta_3 = 0$. We refer the LM test for this hypothesis as LM_1.

It is straightforward to rework the case of the (possibly) asymmetric DLSTAR model. We take a third-order Taylor series approximation of $G(z_t;\gamma_1,\gamma_2,c) = (1 + \exp\{-\gamma_1(z_t - c)\})^{-1} + (1 + \exp\{\gamma_2(z_t + c)\})^{-1}$ with respect to γ_1 and γ_2 evaluated at $\gamma_1 = \gamma_2 = 0$. Since the second derivative is zero, the derived expansion is $G(z_t;\gamma_1,\gamma_2,c) \approx \left[\frac{\gamma_1}{4}(x_t - c) - \frac{\gamma_1^3}{48}(x_t - c)^3\right] + \left[-\frac{\gamma_2}{4}(x_t + c) + \frac{\gamma_2^3}{48}(x_t + c)^3\right]$ yielding the following auxiliary regression:

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t\beta_2 + x_t^{'}z_t^2\beta_3 + x_t^{'}z_t^3\beta_4 + e_t$$

when $\gamma_1 \neq \gamma_2$. Here, the linearity test against the asymmetric DLSTAR model is identical to test the joint restriction that all nonlinear terms are zero: $\beta_2 = \beta_3 = \beta_4 = 0$. We refer the LM test for this hypothesis as LM_2.

When we consider the symmetric case $\gamma = \gamma_1 = \gamma_2$, the derived expansion has a simpler form $G(z_t;\gamma,\gamma,c) \approx \left(\gamma^3 c/24 - \gamma c/2\right) + (\gamma^3 c/8)z_t^2$ yielding the following auxiliary regression:

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t^2\beta_3 + e_t$$

The linearity test against the symmetric DLSTAR model is identical to the joint hypothesis that all nonlinear terms are zero: $\beta_3 = 0$.

A.2. Specification test among ESTAR, symmetric DLSTAR and asymmetric DLSTAR models

Teräsvirta (1994) proposed a specification test of the LSTAR model with a single logistic transition function against ESTAR model based on the auxiliary regression. Similarly, we can use the auxiliary regression equation for the asymmetric DLSTAR model ($\gamma_1 \neq \gamma_2$)

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t\beta_2 + x_t^{'}z_t^2\beta_3 + x_t^{'}z_t^3\beta_4 + e_t$$

which nests that for the symmetric DLSTAR model ($\gamma_1 = \gamma_2$)

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t^2\beta_3 + e_t.$$

The test for the symmetric DLSTAR model against the asymmetric DLSTAR model is identical to test the joint restriction $\beta_2 = \beta_4 = 0$ in the auxiliary regression equation for the asymmetric DLSTAR. We refer the LM test for this hypothesis as LM_3.

Since the auxiliary regression equation for the ESTAR model is

$$y_t = x_t^{'}\beta_1 + x_t^{'}z_t\beta_2 + x_t^{'}z_t^2\beta_3 + e_t$$

the test for the ESTAR model against the asymmetric DLSTAR model ($\gamma_1 \neq \gamma_2$) is identical to test the joint restriction $\beta_4 = 0$ in the auxiliary regression equation for the asymmetric DLSTAR model. We refer the LM test for this hypothesis as LM_4.

Finally, to test for the symmetric DLSTAR model ($\gamma_1 = \gamma_2$) against the ESTAR model, we test the joint restriction $\beta_2 = 0$ in the auxiliary regression equation for the ESTAR model. We refer the LM test for this hypothesis as LM_5.

References

[1] Ball, L., Mankiw, N. G., 1994. Asymmetric price adjustment and economic fluctuations. Economic Journal 104, 247-261.

[2] Ball, L., Mankiw, N. G., Romer, D., 1988. The New Keynesian economics and the output-inflation trade-off. Brookings Papers on Economic Activity 1, 1-65.

[3] Bec, F., Ben Salem, M., Carrasco, M., 2004. Detecting mean reversion in real exchange rates from a multiple regime STAR model. Mimeo.

[4] Calvo, G. A, 1983. Staggered prices in a utility-maximizing framework. Journal of Monetary Economics 12, 383-393.

[5] Calvo, G. A., Reinhart, C. M., 2002. Fear of floating. Quarterly Journal of Economics 117(2), 379-408.

[6] Campa, J. M., Goldberg, L. S., 2005. Exchange rate pass-through to domestic prices. Review of Economics and Statistics 87(4), 679-690.

[7] Choudhri, E. U., Hakura, D. S., 2006. Exchange rate pass-through to domestic prices: Does the inflationary environment matter? Journal of International Money and Finance 25(4), 614-639.

[8] Choudhri, E. U., Faruqee, H., Hakura, D. S., 2005. Explaining the exchange rate pass-through in different prices. Journal of International Economics 65(2), 349-374.

[9] Christiano, L. J., Eichenbaum, M., Evans, C.L., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of Political Economy 113, 1-45.

[10] Devereux, M. B., Siu, H. E., 2007. State dependent pricing and business cycle asymmetries. International Economic Review 48(1), 281-310.

[11] Devereux, M. B., Yetman, J., 2002. Menu costs and the long-run output-inflation trade-off. Economics Letters, 76, 95-100.

[12] Devereux, M. B., Yetman, J., 2008. Price adjustment and exchange pass-through. Journal of International Money and Finance, forthcoming.

[13] Engel, C., 2006. Equivalence results for optimal pass-through, optimal indexing to exchange rates, and optimal choice of currency for export pricing. Journal of the European Economic Association 4(6), 1249-1260.

[14] Goldberg, L. S., Campa, J. M., 2008. The sensitivity of the CPI to exchange rates: Distribution margins, imported inputs, and trade exposure. Review of Economics and Statistics, forthcoming.

[15] Goldberg, P.K., Knetter, M. M., 1997. Goods prices and exchange rates: What have we learned? Journal of Economic Literature 35, 1243-1272.

[16] Gopinath, G., Itskhoki, O., Rigobon, R., 2007. Currency choice and exchange rate pass-through. NBER Working Paper No. W13432.

[17] Granger, C. W. J., Teräsvirta, T., 1993. Modelling Nonlinear Economic Relationships. Oxford University Press, Oxford.

[18] Haggan, V., Ozaki, T., 1981. Modelling nonlinear random vibrations using an amplitude-dependent autoregressive time series model. Biometrika 68(2), 189-196.

[19] Herzberg, V., Kapetanios, G., Price, S., 2003. Import prices and exchange rate pass-through: Theory and evidence from the United Kingdom. Bank of England Working Paper, No. 182.

[20] Kilian, L., Taylor, M. P., 2003. Why is it so difficult to beat the random walk forecast of exchange rates? Journal of International Economics 60, 85-107.

[21] McCarthy, J., 2007. Pass-through of exchange rates and import prices to domestic inflation in some industrialized economies. Eastern Economic Journal 33(4), 511-537.

[22] Michael, P., Nobay, A. R., Peel, D. A., 1997. Transaction costs and nonlinear adjustments in real exchange rates: An empirical investigation. Journal of Political Economy 105, 862-879.

[23] Otani, A., Shiratsuka, S., Shirota, T., 2003. The decline in the exchange rate pass-through: Evidence from Japanese import prices. Bank of Japan Monetary and Economic Studies 21(3), 53-81.

[24] Romer, D., 1990, Staggered price setting with endogenous frequency of adjustment. Economics Letters 32, 205-210.

[25] Saikkonen, P., Luukkonen, R., 1988. Lagrange multiplier tests for testing nonlinearities in time series models. Scandinavian Journal of Statistics 15(1) 55-68.

[26] Sekine, T., 2006. Time-varying exchange rate pass-through: Experiences of some industrial countries. BIS Working Paper, No. 202.

[27] Taylor, J. B., 1980. Aggregate dynamics and staggered contracts. Journal of Political Economy 88, 1-23.

[28] Taylor, J. B., 2000. Low inflation, pass-through, and the pricing power of firms. European Economic Review 44, 1389-1408.

[29] Taylor, M. P., Peel, D. A., 2000. Nonlinear adjustment, long-run equilibrium and exchange rate fundamentals. Journal of International Money and Finance 19, 33-53.

[30] Taylor, M. P., Peel, D. A., Sarno, L., 2001. Nonlinear mean-reversion in real exchange rates: toward a solution to the purchasing power parity puzzles. International Economic Review 42, 1015-1042.

[31] Teräsvirta, T., 1994. Specification, estimation, and evaluation of smooth transition autoregressive models. Journal of the American Statistical Association 89, 208-.218.

[32] van Dijk, D., Teräsvirta, T., Franses, P., 2002. Smooth transition autoregressive models: a survey of recent developments. Econometrics Reviews 21, 1-47.

[33] Walsh, C. E., 2003. Monetary Theory and Policy. Second Edition. MIT Press, Cambridge.

Table 1
LM-type tests for STAR nonlinearity

		Transition Variable ($z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$)					
H_0 vs. H_1	Test Statistics	$d = 1$	$d = 2$	$d = 3$	$d = 4$	$d = 5$	$d = 6$
Linear vs. ESTAR	LM_1	4.93	4.12	3.89	3.58	3.03	2.36
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
	LM_1^*	347.1	341.1	341.5	344.6	346.9	344.9
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Linear vs.	LM_2	5.14	4.04	3.57	2.95	2.68	1.98
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
asymmetric DLSTAR	LM_2^*	351.8	355.3	354.1	357.7	358.2	358.4
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Note: Lag length is $N = 13$. We refer the linearity test against the ESTAR model to LM_1 while the linearity test against the DLSTAR to LM_2. In addition, the heteroskedasticity-robust variants of the LM_1 and LM_2 are denoted as LM_1^* and LM_2^*, respectively. (See Granger and Teräsvirta, 1993). Figures in parentheses below LM statistics denote their p-values.

Table 2
Nonlinear least squares estimation results

A: ESTAR parameter estimates

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

$$G\left(z_t; \widehat{\gamma}\right) = 1 - \exp\left\{\underset{(3.017)}{-0.160}\left(\frac{1}{3}\sum_{j=1}^{3}\pi_{t-j}\right)^2 /0.477\right\}$$

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

B: Symmetric DLSTAR parameter estimates

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

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

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

Note: R^2 denotes the coefficient of determination, se is the standard error of the regression, and $LM(1)$ and $LM(1\text{-}12)$ are Lagrange multiplier test statistics for first-order, and up to twelfth-order serial correlations in the residuals, respectively. The LM statistics in the brackets are p-values. Figures in parentheses below coefficients estimates denote their t-statistics.

Table 3
Asymmetric adjustment

Asymmetric DLSTAR parameter estimates

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

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

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

Note: R^2 denotes the coefficient of determination, se is the standard error of the regression, and $LM(1)$ and $LM(\text{1-12})$ are Lagrange multiplier test statistics for first-order, and up to twelfth-order serial correlations in the residuals, respectively. The LM statistics in the brackets are p-values. Figures in parentheses below coefficients estimates denote their t-statistics.

Table 4
LM-type tests for STAR model selection

		Transition Variable ($z_t = d^{-1} \sum_{j=1}^{d} \pi_{t-j}$)					
H_0 vs. H_1	Test Statistics	$d=1$	$d=2$	$d=3$	$d=4$	$d=5$	$d=6$
Symmetric DLSTAR	LM_3	3.86	3.42	2.84	2.47	2.17	1.26
		(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.19)
vs. Asymmetric DLSTAR	LM_3^*	358.2	348.6	350.9	356.2	360.3	361.3
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
ESTAR	LM_4	4.41	3.24	2.52	1.56	1.81	1.19
		(0.00)	(0.00)	(0.01)	(0.10)	(0.05)	(0.29)
vs. Asymmetric DLSTAR	LM_4^*	374.1	376.1	371.3	371.8	372.4	369.2
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Symmetric DLSTAR	LM_5	2.97	3.34	3.02	3.30	2.47	1.32
		(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.20)
vs. ESTAR	LM_5^*	377.0	377.0	373.6	373.1	372.6	369.8
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Note: Lag length is $N = 13$. LM_3 is to test for the symmetric DLSTAR model against the asymmetric DLSTAR model, LM_4 is to test for the ESTAR model against the asymmetric DLSTAR model, LM_5 is to test for the symmetric DLSTAR model against the ESTAR model. In addition, the heteroskedasticity-robust variants of the LM_3, LM_4, LM_5 are denoted as LM_3^*, LM_4^*, LM_5^*, respectively. (See Granger and Teräsvirta, 1993). Figures in parentheses below LM statistics denote their p-values.

Table 5
Demand shocks

ESTAR estimates with industrial production index

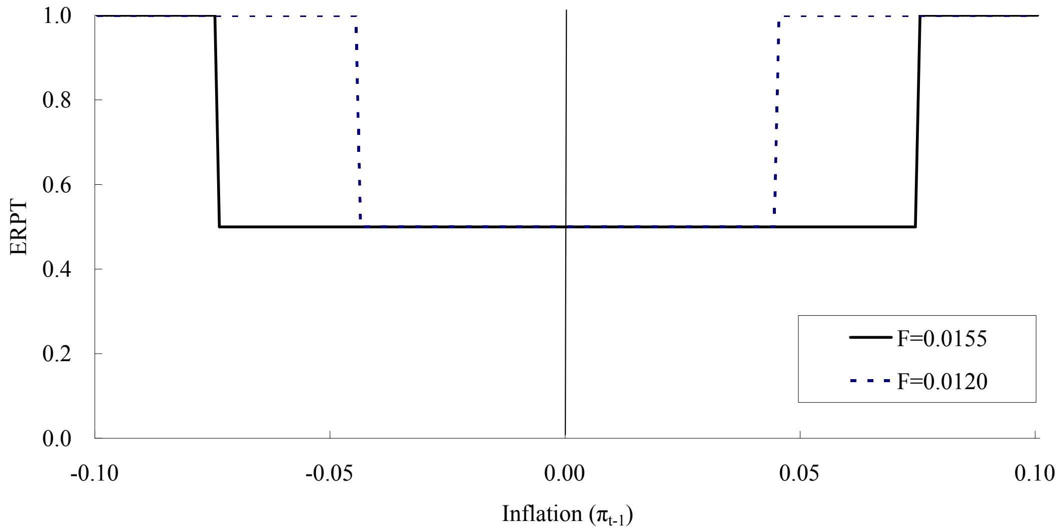
$$
\begin{aligned}
\pi_t \quad = \quad & \underset{(4.107)}{0.119} + \underset{(2.799)}{0.157}\pi_{t-1} + \underset{(5.026)}{0.279}\pi_{t-3} - \underset{(-2.763)}{0.105}\,\pi_{t-5} + \underset{(11.283)}{0.331}\,\Delta(s_t + p_t^*) \\
& + \underset{(2.496)}{0.080}\Delta(s_{t-1} + p_{t-1}^*) - \underset{(-1.052)}{0.042}\,\Delta(s_{t-3} + p_{t-3}^*) + \underset{(3.149)}{0.105}\Delta(s_{t-4} + p_{t-4}^*) - \underset{(-3.252)}{0.126}\,\Delta iip_t + \\
& \Big[- \underset{(-3.528)}{0.865}\,\pi_{t-1} - \underset{(-2.751)}{0.957}\,\pi_{t-3} + \underset{(22.727)}{0.669}\,\Delta(s_t + p_t^*) + \underset{(1.980)}{0.652}\Delta(s_{t-3} + p_{t-3}^*) \\
& - \underset{(-4.080)}{1.312}\,\Delta(s_{t-4} + p_{t-4}^*) + \underset{(2.355)}{0.674}\Delta iip_t - \underset{(-2.221)}{0.717}\,\Delta iip_{t-5} \Big] \, G\left(z_t; \widehat{\gamma}\right) + \widehat{\varepsilon}_t,
\end{aligned}
$$

$$
G\left(z_t; \widehat{\gamma}\right) = 1 - \exp\left\{ \underset{(4.010)}{-0.100} \left(\frac{1}{2} \sum_{j=2}^{3} \pi_{t-j} \right)^2 / 0.551 \right\}
$$

$R^2 = 0.623,\ se = 0.467,\ obs = 396,\ LM(1) = [0.145],\ LM(\text{1-12}) = [0.024]$

Note: R^2 denotes the coefficients of determination, se is the standard error of the regression, and $LM(1)$ and $LM(\text{1-12})$ are Lagrange multiplier test statistics for first-order, and up to twelfth-order serial correlations in the residuals, respectively. The LM statistics in the brackets are p-values. Figures in parentheses below coefficients estimates denote their t-statistics.

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



Note: $\sigma^2 = 0.01$.

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

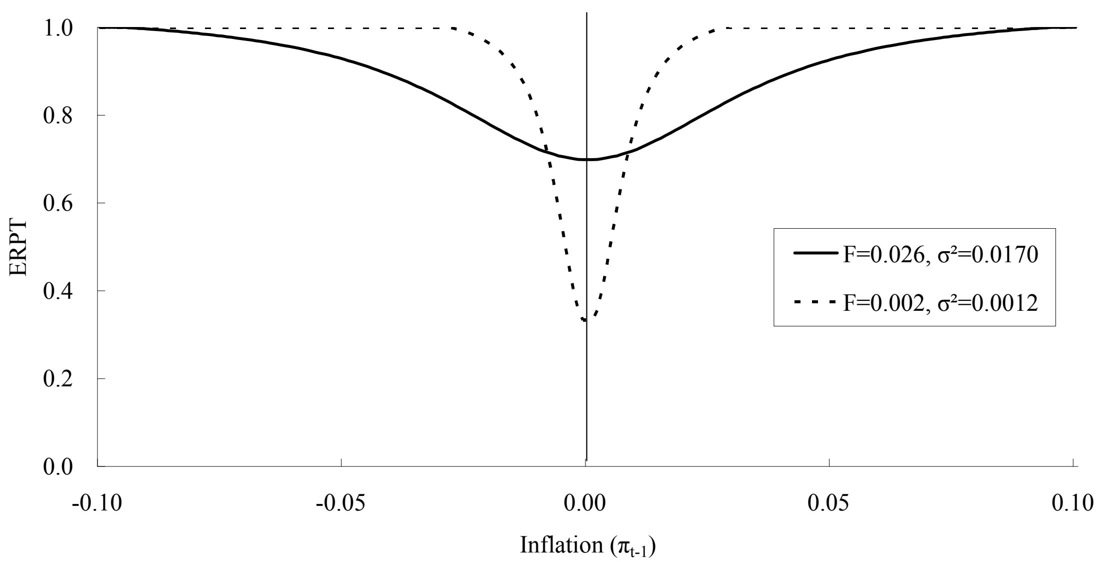
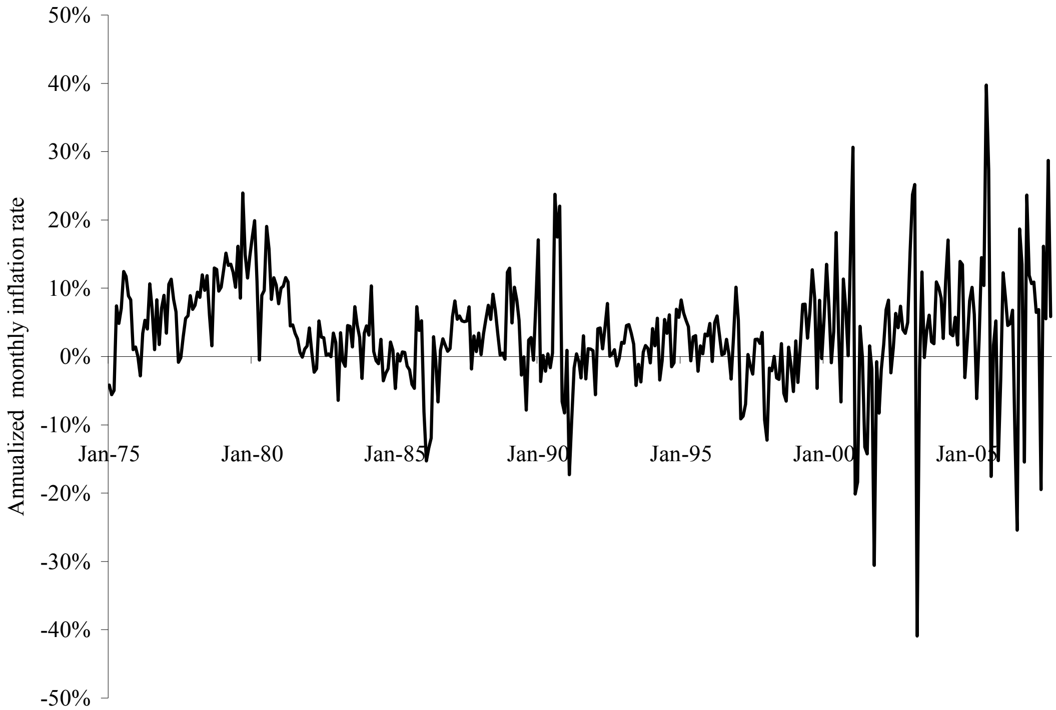


Figure 3. Producer price index inflation



Note: Seasonally adjusted series.

Figure 4. ERPT against transition variable: ESTAR model

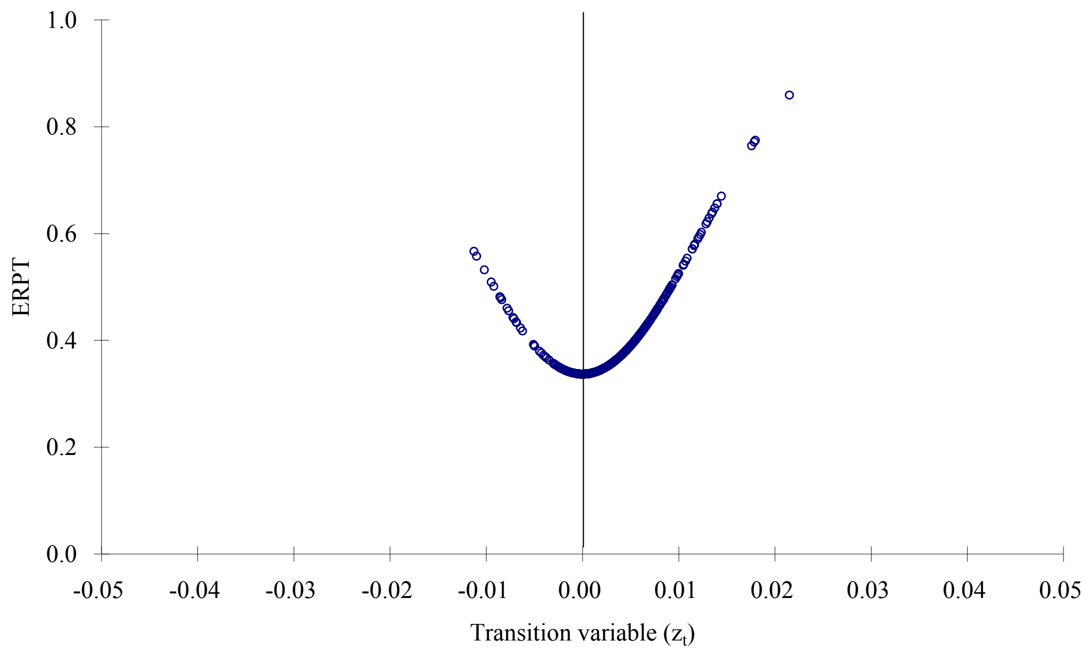


Figure 5. ERPT over time: ESTAR model

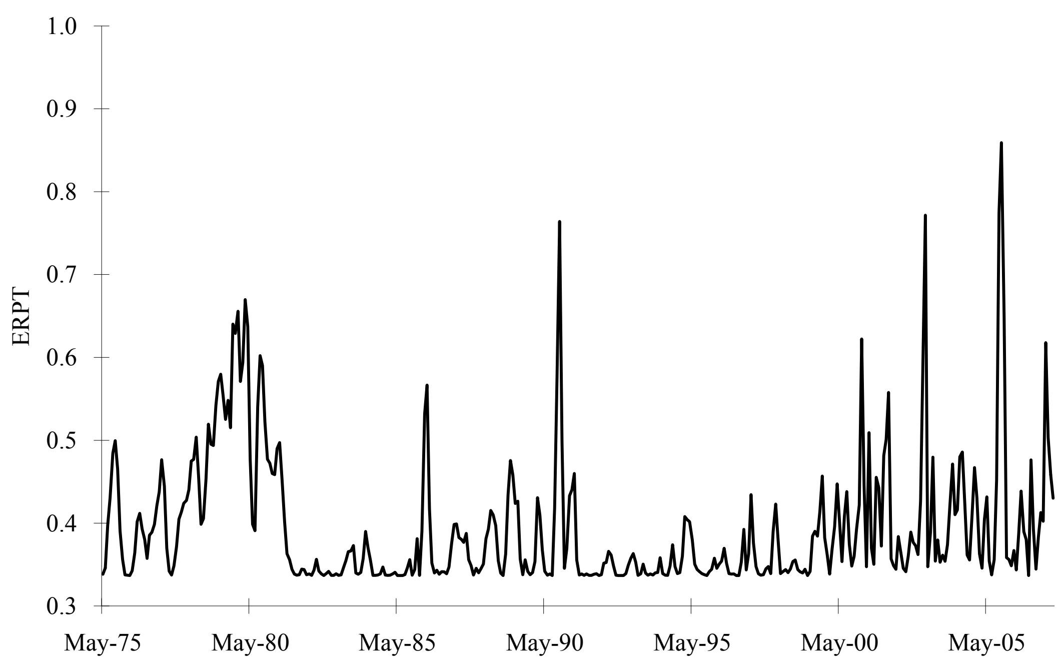


Figure 6. ERPT against transition variable: Symmetric DLSTAR model

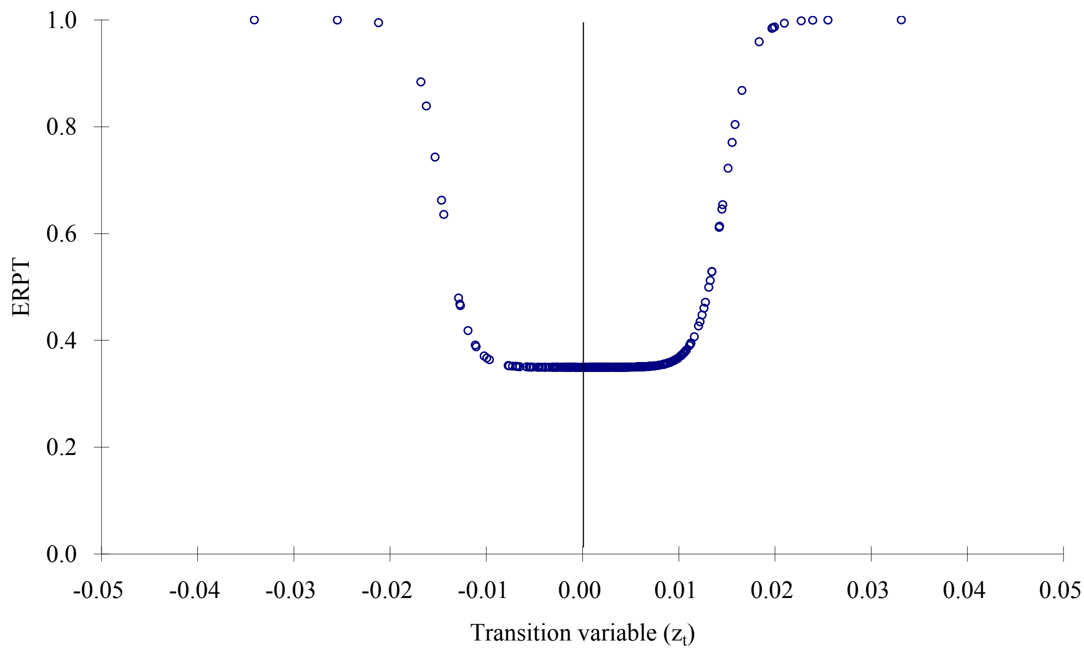


Figure 7. ERPT over time: Symmetric DLSTAR model

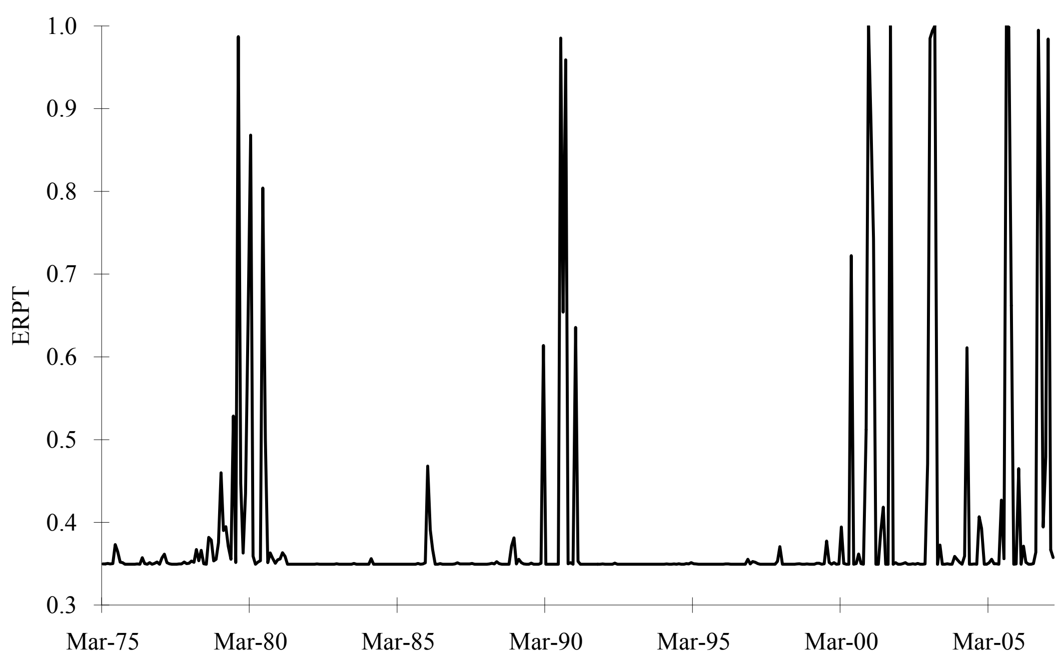


Figure 8. ERPT against transition variable: Asymmetric DLSTAR model

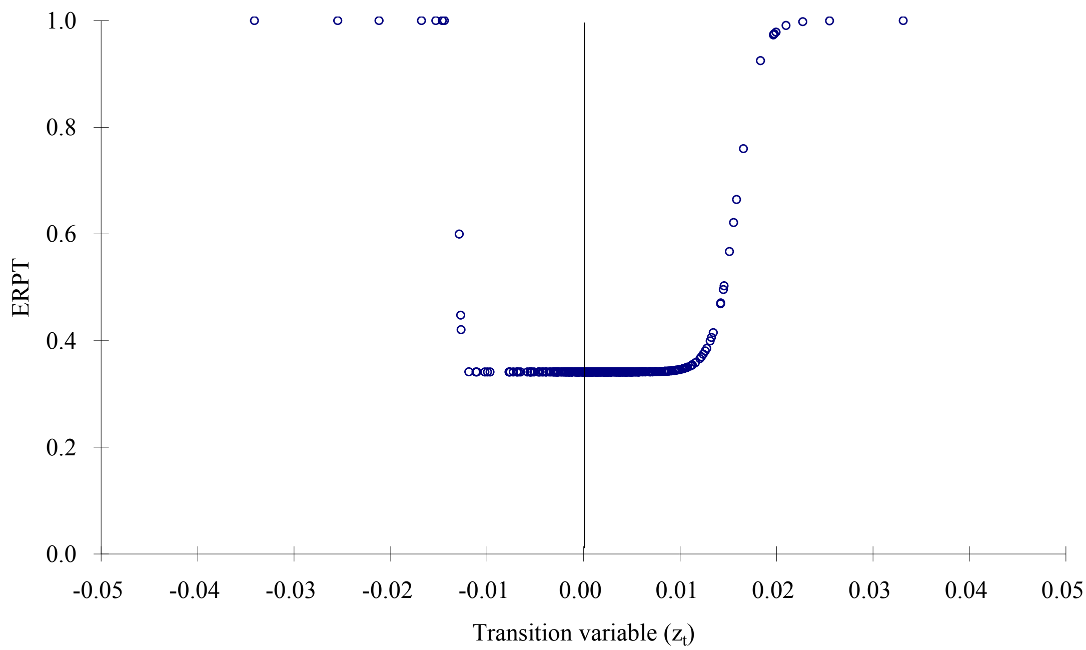


Figure 9. ERPT over time: Asymmetric DLSTAR model

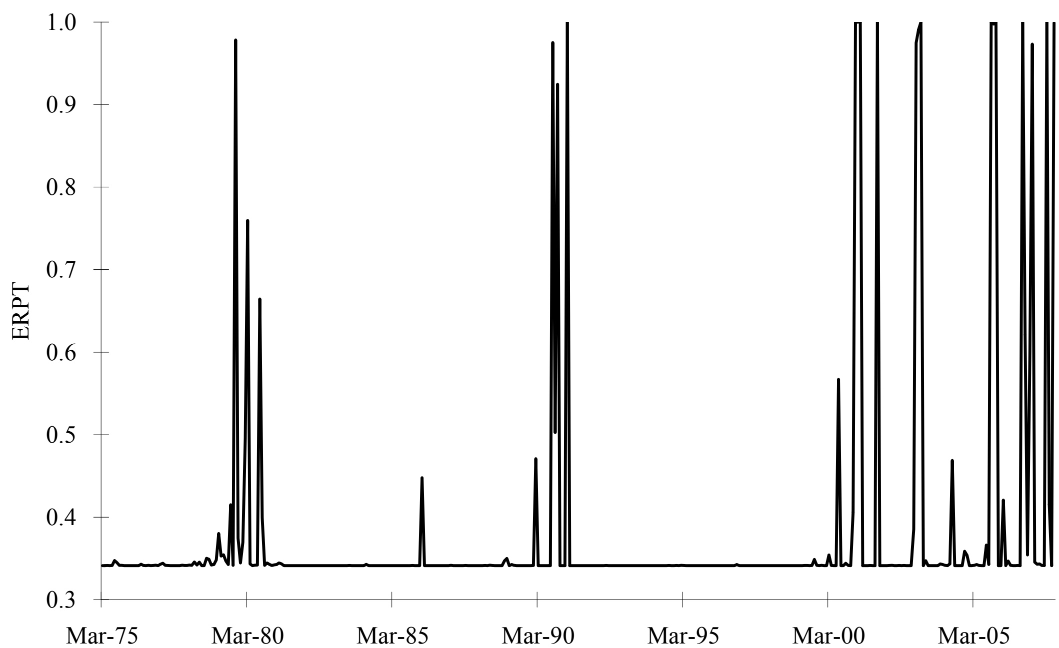


Figure 10. ERPT over time: ESTAR model with industrial production index