Exchange Rate Pass-Through to Consumer Prices and the Role of Energy Prices

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Abstract

A group of researchers has asserted that the rate of exchange rate pass-through (ERPT) to domestic prices has declined substantially over the last few decades. We revisit this claim of a downward trend in the rate of ERPT to the Consumer Price Index (CPI) by employing the vector autoregressive (VAR) model for the U.S. macroeconomic data under the current floating exchange rate regime. Our VAR approach that nests the conventional single equation method reveals very weak evidence of ERPT during the pre-1990 era. On the other hand, we observe statistically significant evidence of ERPT during the post-1990 era, which sharply contrasts with previous findings. After statistically confirming a structural break in ERPT to the total CPI via Hansen's (2001) test procedure, we seek the source of the structural break using the disaggregate level CPIs, which pinned down a key role of energy prices in explaining the emergence of the break. The dependency of the U.S. energy consumption on imports has increased since the 1990s. This change magnifies the effects of the exchange rate shock on domestic energy prices, resulting in greater responses of the total CPI via this energy price channel.

Keywords: Exchange Rate Pass Through; Disaggregated CPI; Structural Break; Oil Price Shock

JEL Classification: E31; F31; F41

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1 Introduction

In an open economy, changes in the value of the home currency can substantially influence the Consumer Price Index (CPI) of the home country. In response to depreciations of the home currency, import goods prices may rise, triggering increases in the CPI of the home country. In addition, the relative demand for domestic products may grow through the expenditure spending effect when the home currency loses value against the foreign currency, making the home products become relative cheaper. Stronger demand for domestic goods then raises home consumer prices.

Rates of exchange rate pass-through (ERPT) to the CPI may change over time. Frankel, Parsley, and Wei (2012), for example, provide international evidence that the degree of ERPT to import goods prices, and in turn to the CPI, has decreased in the 1990s, especially more substantially in developing countries. Campa and Goldberg (2005) report similar evidence for 23 high-income OECD countries. Taylor (2000) claims that a downward trend in ERPT to the U.S. CPI in the 1990s is due to a lower inflationary environment. Gagnon and Ihrig (2004) extended Taylor's claim to 11 industrialized countries. Similarly, Takhtamanova (2010) presents empirical evidence of structural breaks in the rate of ERPT during the 1990s for a set of fourteen OECD countries.

We revisit this claim of a downward trend in the rate of ERPT to domestic prices, employing the structural vector autoregressive (VAR) model for the U.S. macroeconomic data under the current floating exchange rate regime since 1973. Our empirical findings sharply contrast with aforementioned earlier findings that heavily rely on single equation approaches.

As we can see in Figure 1, we obtained statistically significant negative responses of the total CPI to the real exchange rate shock only when the post-1990 (1990:I-2017:IV) sample period is used. When we employ the pre-1990 (1973:II-1989:IV) sample period, the total CPI responds initially positively then rapidly declines below zero. Furthermore, it comes with a very wide 90% confidence band. These qualitatively different responses imply the presence of a structural break in the rate of ERPT to the CPI in the U.S.

Figure 1 around here

We statistically test this possibility by employing Hansen's (2001) sequential structural break test procedure. We identified a structural break in the rate of ERPT to an array of disaggregated CPIs including the total CPI that occur roughly at the beginning of the 1990s. We also estimate and report the ERPT rate parameters from the structural VAR method that nests the conventional single equation approach. We report statistically significant evidence of ERPT to the CPIs only for the post-1990 era.

Then, we seek the source of the structural break by looking at the disaggregated CPI responses to the real exchange rate shock using the impulse-response function analysis. We demonstrate that the time-varying responses of the total CPI are driven mainly by the responses of the energy-related CPIs. That is, we show that energy prices play an important role in explaining the existence of the structural break. We note that the dependency of U.S. oil consumption on imports substantially increased in the post-1990 era, which may explain the presence of the structural break if there exists a strong link between the oil price and the CPI. We confirm this conjecture using a quad-VAR model that uncovers a strong oil price channel that propagates the exchange rate shock to the CPI.

The remainder of this paper is organized as follows. Section 2 presents our baseline empirical model. We also explain how our multivariate VAR approach for ERPT nests the conventional single equation method. Section 3 reports structural break test results. In Section 4, we implement the impulse-response function analysis that unveils an important role of energy prices in explaining the source of the break. Section 5 concludes.

2 The Empirical Model

We employ a recursively identified vector autoregressive (VAR) process of order q to study the dynamic exchange rate effects on consumer prices in the U.S. Abstracting from deterministic terms, we propose the following model.[1]

$$\mathbf{x}_t = \sum_{j=1}^{q} \mathbf{B}_j \mathbf{x}_{t-j} + \mathbf{C}\mathbf{u}_t, \tag{1}$$

where

$$\mathbf{x}_t = [\Delta s_t \quad \Delta y_t \quad \Delta p_t]',$$

$\mathbf{C}$ denotes a lower-triangular (Choleski factor) matrix, and $\mathbf{u}_t$ is a vector of mutually orthonormal structural shocks, that is, $E\mathbf{u}_t\mathbf{u}_t^{'} = \mathbf{I}$. s_t denotes the real exchange rate, y_t is the U.S. real GDP, and p_t is the U.S. Consumer Price Index (CPI). All variables are log transformed and differenced, so positive values of Δs_t, for example, denote real appreciations of the U.S. dollar.

We are particularly interested in the j-period ahead orthogonalized impulse-response function (IRF) of domestic CPI inflation (Δp_t) to the structural shock to Δs_t that occurs at time

[1] We demean the VAR prior to estimations.

t. That is,

$$\phi(j) = E\left(\Delta p_{t+j} | u_{\Delta s,t} = 1, \Omega_{t-1}\right) - E\left(\Delta p_{t+j} | \Omega_{t-1}\right), \tag{2}$$

where $E(\cdot|\Omega_{t-1})$ is the conditional expectation operator given the information set Ω_{t-1} at time $t-1$, and $u_{\Delta s,t}$ is the 1% shock to Δs_t at time t.[2] The j-period ahead *level* CPI response is obtained by the following cumulative sum.

$$\eta(j) = \sum_{s=0}^{j} \phi(s) \tag{3}$$

It is well known that econometric inferences from recursively identified VAR models such as (1) may not be robust to alternative VAR ordering. It should be noted that our empirical analysis doesn't suffer from this criticism given the fixed location of Δs_t in $\mathbf{x}_t$ as long as we are interested only in exchange rate pass-through (ERPT) to consumer prices. That is, all IRF estimates to the exchange rate (Δs_t) shock, $\phi(j)$ and $\eta(j)$, are *numerically identical* even if one randomly re-shuffles the variables next to Δs_t in (1). See Christiano, Eichenbaum, and Evans (1999) for details. Therefore, our key findings presented in this paper are robust to alternative ordering.

One may estimate $\boldsymbol{\alpha} = [\alpha_{s,0}...\alpha_{s,q}\ \alpha_{y,0}...\alpha_{y,q}\ \alpha_{p,1}...\alpha_{p,q}]'$ from the following univariate equation to measure the degree of ERPT. See Takhtamanova (2010) and Campa and Goldberg (2005), among others.[3]

$$\Delta p_t = \sum_{j=0}^{q} \alpha_{s,j} \Delta s_{t-j} + \sum_{j=0}^{q} \alpha_{y,j} \Delta y_{t-j} + \sum_{j=1}^{q} \alpha_{p,j} \Delta p_{t-j} + u_t \tag{4}$$

Then, the long-run measure of ERPT is defined as,

$$ERPT = \frac{\sum_{j=0}^{q} \alpha_{s,j}}{1 - \sum_{j=1}^{q} \alpha_{p,j}}, \tag{5}$$

while $\alpha_{s,0}$ in (4) provides information on the short-run ERPT.

It should be noted that the VAR model in (1) nests this conventional measure of ERPT in (5). Obtaining the Choleski factor ($\mathbf{C}$) from the reduced form estimation of (1), we pre-

[2]The information set Ω_t is adaptive in the sense that $\Omega_j \supseteq \Omega_{j-1}, \forall j$.

[3]Additional control variables can be added in the right hand side of the equation, which can be also easily nested by VAR models following the procedure in the present paper.

multiply each side of (1) by $\mathbf{C}^{-1}$ to recover the structural form VAR:

$$\mathbf{A}_0\mathbf{x}_t = \sum_{j=1}^{q} \mathbf{A}_j\mathbf{x}_{t-j} + \mathbf{u}_t, \tag{6}$$

where $\mathbf{A}_0 = \mathbf{C}^{-1}$ and $\mathbf{A}_j = \mathbf{C}^{-1}\mathbf{B}_j$. By dividing the third equation of the VAR system (6) by the $(3,3)$ component of $\mathbf{A}_0$, we obtain the following *normalized* equation which is equivalent to (4).

$$\Delta p_t = \sum_{j=0}^{q} \beta_{s,j}\Delta s_{t-j} + \sum_{j=0}^{q} \beta_{y,j}\Delta y_{t-j} + \sum_{j=1}^{q} \beta_{p,j}\Delta p_{t-j} + \varepsilon_t \tag{7}$$

Note that (7) parameter estimates for $\boldsymbol{\beta} = \left[\beta_{s,0}...\beta_{s,q}\ \beta_{y,0}...\beta_{y,q}\ \beta_{p,1}...\beta_{p,q}\right]'$ are not the same as $\boldsymbol{\alpha}$ parameters in (4) in finite samples, because $\boldsymbol{\alpha}$ is estimated via the least squares regression for (4), while $\boldsymbol{\beta}$ is estimated from the VAR model (1). In addition to the conventional measure of ERPT, $\frac{\sum_{j=0}^{q} \beta_{s,j}}{1-\sum_{j=1}^{q} \beta_{p,j}}$, our approach also allows dynamic measures of ERPT at any point of time horizon j, which can be obtained via the impulse-response function $\eta(j)$.

3 The Empirics

3.1 Data Descriptions

We obtained most data from the Federal Reserve Economic Data (FRED) website. Observations span from 1973:I to 2017:IV for the post-Bretton Woods system. That is, we focus on exchange rate pass-through (ERPT) during the current floating exchange rate regime.

y_t is the U.S. real Gross Domestic Product (GDPC1). s_t denotes the real trade weighted U.S. dollar index with major currencies (TWEXMPA). p_t is the Consumer Price Index. In addition to the total CPI (CPIAUCSL), we consider an array of disaggregated consumer price sub-indices that includes: Food CPI (CPIUFDSL); Housing CPI (CPIHOSSL); Apparel CPI (CPIAPPSL); Transportation CPI (CPITRNSL); Medical Care CPI (CPIMEDSL); Energy CPI (CPIENGSL); All Items less Energy CPI (CPILEGSL); All Items less Food CPI (CPIULFSL); All Items less Food and Energy CPI (CPILFESL). In what follows, we use these disaggregate level CPIs to search for the source of a structural break in the rate of ERPT to the total CPI.

We also use the spot crude oil price (West Texas Intermediate: WTISPLC) in our extended VAR models. We deflated this nominal crude oil price by the U.S. CPI to get the real oil price rop_t. We transformed monthly frequency data (s_t, p_t, rop_t) to quarterly frequency ones by taking end of period values to match the quarterly frequency of y_t. All data are log differenced to ensure the stationarity of the VAR model (1).

3.2 Structural Break in Exchange Rate Pass-Through to CPI Inflation

To statistically evaluate the conjecture on the existence of a structural break, we implement an array of econometric tests for (7), employing the sequential test procedure suggested by Hansen (2001). Let $\boldsymbol{\beta}^R = \left[\beta^R_{s,0}...\beta^R_{s,q}\ \beta^R_{y,0}...\beta^R_{y,q}\ \beta^R_{p,1}...\beta^R_{p,q}\right]'$, $R = 1, 2$ for the following stochastic process.

$$\Delta p_t = \sum_{j=0}^{q} \beta^R_{s,j}\Delta s_{t-j} + \sum_{j=0}^{q} \beta^R_{y,j}\Delta y_{t-j} + \sum_{j=1}^{q} \beta^R_{p,j}\Delta p_{t-j} + \varepsilon_t,\ R = 1, 2, \tag{8}$$

where

$$\begin{aligned} \boldsymbol{\beta} &= \boldsymbol{\beta}^1,\ t = \left[0, \tau^1\right) \\ \boldsymbol{\beta} &= \boldsymbol{\beta}^2,\ t = \left[\tau^1, T\right], \end{aligned}$$

under the alternative hypothesis H_A.

We begin the procedure by testing the null hypothesis $H_0 : \boldsymbol{\beta}^1 = \boldsymbol{\beta}^2$, employing the following three test statistics proposed by Andrews (1993) and Andrews and Ploberger (1994) for the full sample (T).

$$\begin{aligned} \text{Sup}F_T &= \sup_{k_1 \le k \le k_2} F_T(k) \\ \text{Exp}F_T &= \ln\left(\frac{1}{k_2 - k_1 + 1}\sum_{t=k_1}^{k_2} \exp\left(\frac{1}{2}F_T(k)\right)\right) \\ \text{Ave}F_T &= \frac{1}{k_2 - k_1 + 1}\sum_{t=k_1}^{k_2} F_T(k), \end{aligned} \tag{9}$$

where $F_T(k)$ denotes the Lagrange multiplier (LM) test statistics of the null hypothesis of no structural change at each grid point $k \in [k_1, k_2]$.[4]. We use conventional values for the trimming parameters, $k_1 = 0.15T$, $k_2 = 0.85T$. Associated p values are obtained by numerical approximations to the asymptotic distributions of the test statistics using the method by Hansen (1997).

When the test rejects the null hypothesis of no structural break at time $t = \tau^1$, we repeat the test for the resulting two sub-sample periods, $[0, \tau^1)$ and $\left[\tau^1, T\right]$, for the existence of

[4] Alternatively, the Wald or the Likelihood Ratio statistics can be used.

additional structural break date, say $t = \tau^2$, in each sub-sample. If the test identifies a break in $[0, \tau^1)$, we implement the test for the two sub-sample periods, $[0, \tau^2)$ and $\left[\tau^2, \tau^1\right]$ for another structural break at time $t = \tau^3$ in each sub-periods. If the test fails to identify a break in $[0, \tau^1)$, we stop looking for break dates in that region. We repeat this for the other sub-sample period $\left[\tau^1, T\right]$ until the test fails to reject the null hypothesis. We choose the number of lags by the Akaike Information Criteria (AIC) with a maximum 4 lags.

As can be seen in Table 1, we obtained strong evidence of a structural break from the full sample period for all CPIs with exceptions of the CPIs that exclude energy prices, that is, the All less Energy CPI and the All less Food and Energy CPI. For example, the null hypothesis of no structural break was rejected by all three tests (9) for the All item CPI at the 1% significance level, while all three tests fail to reject the null hypothesis even at the 10% significance level for the All less Food and Energy CPI.

We report the identified break date from the SupF_T test in Table 1, which selects the break date around the late 1980s whenever the test rejects the null hypothesis. The test fails to reject the null hypothesis for the sub-sample periods implied by the chosen break date from the first test at any conventional significance level.[5] That is, our structural break tests identified a single statistically significant break date for most CPIs other than the ones that exclude energy prices. These findings imply that the structural break in the total CPI may be mainly due to the break in the Energy CPI.

Table 1 around here

4 Searching for the Source of the Break

4.1 Time-Varying Degree of ERPT to CPI

Confirming the statistical evidence of a structural break in ERPT to U.S. CPIs, we now seek the source of the structural break by investigating disaggregate level CPI responses to the real exchange rate shock via the impulse-response function analysis. As we observed in the previous section, the SupF_T test identified the late 1980s as a common statistically significant single break date for most CPIs. Based on these results, we estimate the degree of ERPT for the two sub-sample periods, the pre-1990 (1973:I-1989:IV) and the post-1990 (1990:I-2017:IV) eras. We report the ERPT parameter estimates from the IRF approach in

[5] These results are available upon request.

Table 2, while the ones from the structural VAR coefficient approach appear in Table 3.[6] Interesting findings are as follows.

First, we obtain statistically insignificant estimates for all CPIs during the pre-1990 era, while estimates from the post-1990 era are highly significant for 6 out of 10 CPIs no matter what approaches were used. These findings imply strong evidence of ERPT to CPIs only in the later sample period. Furthermore, all significant ERPT parameter estimates turn out to be negative, which means that appreciations of the U.S. dollar result in decreases in CPIs as we discussed in the introduction.

Second, we confirm a weaker rate of ERPT to the total CPI during the post-1990 era, which is consistent with the works of Frankel, Parsley, and Wei (2012), Takhtamanova (2010), McCarthy (2007), Campa and Goldberg (2005), Gagnon and Ihrig (2004), and Taylor (2000) who claim that the degree of ERPT became substantially weaker in the 1990s. The ERPT *point* estimate with the total CPI from the VAR coefficient approach (see Table 3) were -0.179 and -0.134 for the pre- and the post-1990 eras, respectively. Although these point estimates are consistent with the previous works, one should note that it is statistically significant only for the post-1990 sample period. We obtain similar results from the IRF approach as can be seen in Table 2. Therefore, our estimation results cast serious doubt on the validity of previous findings because our structural estimation approaches reveal high degree of uncertainty in the pre-1990 era.

Third, ERPT estimates tend to be greater in absolute value for CPIs with food and energy prices during the post-1990 era. ERPT estimates for the Housing, Apparel, and the Medical Care CPIs are not only quantitatively smaller but also statistically insignificant, meaning that the significant and strong rate of ERPT to the total CPI is mainly driven by the strong degree of ERPT to some of its sub-category CPIs such as the Energy CPI, the Transportation CPI, and the Food CPI. It should be noted that the ERPT estimate becomes statistically insignificant and negligibly small when food and energy prices are excluded (All less F&E CPI). Note also that the ERPT estimate of the CPI without Food is greater than that of the CPI without Energy, even though they are both significant, which implies a stronger role of the Energy CPI in comparison with the Food CPI in contributing to ERPT to the total CPI in the post-1990 era.

Tables 2 and 3 around here

[6] We obtain the long-run rate of ERPT from the IRF approach by taking the value of $\eta(20)$. $j = 20$ (quarter) seems to be sufficient for the IRF to get stabilized.

4.2 Time-Varying Responses of Disaggregated CPIs

This subsection further investigates the source of the structural break in ERPT to the total CPI by investigating the IRFs of disaggregated level CPIs to the real exchange rate shock in the pre- and the post-1990 eras.

We note that all IRFs of the CPIs for the pre-1990 era in Figure 2 exhibit positive responses, although insignificant, with an exception of the Food CPI. We obtained weak evidence of ERPT only to the Food CPI during the pre-1990 era, which implies that the negative responses of the total CPI we observed in Figure 1 should have been driven solely by ERPT to the Food CPI given virtually negligible responses of other CPIs. On the other hand, we observed statistically significant negative responses of the Food CPI, Transportation CPI, and the Energy CPI during the post-1990 era. Responses were negligibly weak and overall insignificant for the Housing CPI, the Apparel CPI, and the Medical CPI.

Interestingly, as can be seen in Figure 3, the All less Energy CPI and the All less Food CPI exhibit significantly negative responses during the post-1990 era, but the All less Food and Energy CPI shows no evidence of ERPT in that period. These responses imply that both food and energy prices play an important role in explaining ERPT to the total CPI during the post-1990's era.

It should be also noted that the Energy CPI and its related prices such as the Transportation CPI play a key role in explaining the existence of the structural break in ERPT to the total CPI. This is because ERPT to the Food CPI was observed in both sub-sample periods, whereas we observe more significant evidence only for the post-1990 sample period. Unlike the Food CPI, the Energy CPI responses are qualitatively different across the two sample periods. That is, we observe negative responses of the Energy CPI only in the post-1990 sample period, which resembles the responses of the total CPI to the exchange rate shock. To put it differently, we pinned down the structural change in ERPT to the Energy CPI as the source of the structural break in ERPT to the total CPI in the U.S.

Figures 2 and 3 around here

4.3 Role of Energy Prices in Explaining the Break

The previous section presented an important role of energy prices in explaining the break in ERPT to the total CPI in the U.S. To better understand this phenomenon, we report the share of the U.S. petroleum net imports in Figure 4.[7] The share exhibits an overall positive

[7] We obtained the data from the U.S. Energy Information Administration website.

trend until 1978 followed by decreases until the mid 1980s. Then, it rapidly went up again until the emergence of the U.S. sub-prime mortgage market crisis in 2006, causing the share to rapidly shrink.

We note that the consumption share stays roughly well below 40% most of the time during the pre-1990 era, while the share increased rapidly toward its historical peak of 60% in 2006. Since then, it started declining due to the U.S. financial crisis followed by the Great Recession. Nonetheless, the average ratio in the post-1990 era is 46%, which is substantially higher than the pre-1990 era average 24%. As the energy dependency increases in the U.S. since the 1990s, the effect of the exchange rate shock on domestic energy prices should have become stronger. Given that, a strong link between energy prices and overall domestic prices (total CPI) will explain the emergence of the structural break in the pattern of ERPT to the total CPI via the energy price channel.

Figure 4 around here

To investigate this possibility, we report the IRFs from the following quad-variate VAR system.

$$\mathbf{x}_t = \sum_{j=1}^{q} \mathbf{B}_j \mathbf{x}_{t-j} + \mathbf{C}\mathbf{u}_t, \tag{10}$$

$$\mathbf{x}_t = [\Delta s_t \quad \Delta rop_t \quad \Delta y_t \quad \Delta p_t]',$$

where rop_t is the log real oil price. Results are reported in Figure 5. Note that the location of Δrop_t in $\mathbf{x}_t$ doesn't matter as long as we are interested in the effect of the exchange rate shock on the other variables.

The real oil price responds significantly negatively to the real exchange rate shock during the post-1990 era, while it exhibits insignificant positive responses during the earlier period. The real U.S. GDP responses are overall insignificantly negative in both periods, and are quantitatively negligible. The responses of the total CPI and the Energy CPI are similar with each other and they resemble those of the real oil price, which confirm our claim that the break in the pattern of ERPT to the total CPI is caused by the break in the response of the oil price to the exchange rate shock.

Figure 5 around here

As a robustness check, we also implement the structural break test for the total CPI with a quad-variate VAR model as follows.

$$\Delta p_t = \sum_{j=0}^{q} \beta_{s,j}^{R} \Delta s_{t-j} + \sum_{j=0}^{q} \beta_{o,j}^{R} \Delta rop_{t-j} + \sum_{j=0}^{q} \beta_{y,j}^{R} \Delta y_{t-j} + \sum_{j=1}^{q} \beta_{p,j}^{R} \Delta p_{t-j} + \varepsilon_t, \;\; R = 1, 2, \quad (11)$$

Hansen's (2001) sequential test procedure rejects the null hypothesis again at least at the 8% significance level, and identifies two break dates, 1993:III and 2006:IV. The first one is roughly consistent with our previous findings from the tri-variate model, while the second one corresponds to the beginning of the recent U.S. financial crisis that triggered rapid decreases in the consumption dependency on foreign energy products (see Figure 4). These findings strengthen our claim of an important role of energy prices in explaining time-varying ERPT to the total CPI in the U.S.[8]

5 Concluding Remarks

This paper revisits the claim of a downward trend in the rate of exchange rate pass-through (ERPT) to the CPI by employing the recursively identified vector autoregressive (VAR) model for the U.S. macroeconomic data during the current floating exchange rate regime. Our findings sharply contrast with those from earlier works, to name a few, Frankel, Parsley, and Wei (2012), Takhtamanova (2010), and Campa and Goldberg (2005), that are mostly obtained from the univariate regression approach. We also demonstrate that our VAR approach allows richer statistical analyses but is flexible enough to nest the single equation approach.

Being consistent with earlier work, our findings suggest some evidence of a greater rate of ERPT to the total CPI in the pre-1990 sample period. However, our multivariate VAR approach reveals high uncertainty on the ERPT estimation during that period, which makes statistical inferences on ERPT to be difficult. On the other hand, we report highly significant negative responses of the total CPI to the real exchange rate shock during the post-1990 era. These qualitatively different responses across the two eras imply the possibility of a structural break in the pattern of ERPT to the CPI in the U.S.

Employing Hansen's (2001) test procedure, we confirm the existence of a structural break at the end of the 1980s. Then, we seek the source of the structural break by investigating the impulse-response function (IRF) estimates of the disaggregate level CPIs to the real exchange rate shock. Our empirical results show that energy prices play a key role in explaining the change in the pattern of ERPT to the total CPI. We point out that the energy dependency

[8] Detailed results are available upon requests.

of the U.S. economy on imports became substantially higher since the 1990s. Greater dependency amplifies the effect of the exchange rate shock on domestic energy prices, resulting in greater responses of the total CPI via the energy price channel. That is, we demonstrate that energy prices play a key role in explaining the structural break in ERPT to the total CPI in the U.S.

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Table 1. Structural Break Test Results

CPI	Sample Period	Break Date	SupF_T	ExpF_T	AveF_T	#Lag
All item	1973Q1-2017Q4	1989Q1	41.24	17.51	29.99	4
			(0.01)	(0.00)	(0.00)	
Food	1973Q1-2017Q4	1989Q1	31.53	13.24	22.97	4
			(0.12)	(0.07)	(0.03)	
Housing	1973Q1-2017Q4	1982Q1	28.59	12.19	20.34	3
			(0.08)	(0.04)	(0.02)	
Apparel	1973Q1-2017Q4	1992Q3	32.30	13.20	18.48	2
			(0.00)	(0.00)	(0.00)	
Transportation	1973Q1-2017Q4	1988Q4	26.44	10.62	18.99	2
			(0.03)	(0.02)	(0.00)	
Medical Care	1973Q1-2017Q4	1989Q4	33.20	13.30	21.83	4
			(0.08)	(0.07)	(0.04)	
Energy	1973Q1-2017Q4	1988Q4	21.75	8.73	15.67	2
			(0.14)	(0.08)	(0.02)	
All less Energy	1973Q1-2017Q4	1983Q1	26.77	11.17	19.99	4
			(0.33)	(0.21)	(0.09)	
All less Food	1973Q1-2017Q4	1988Q3	33.52	14.25	23.92	3
			(0.02)	(0.01)	(0.00)	
All less F&E	1973Q1-2017Q4	1983Q1	27.99	11.14	19.16	4
			(0.26)	(0.21)	(0.13)	

Note: We follow the structural break test procedure suggested by Hansen (2001) by applying the tests (Hansen, 1997) sequentially for the sub-sample periods chosen by the break date estimates. Approximate p values were obtained using Hansen's GAUSS codes, and are reported in parentheses. The number of lags was chosen by the Akaike Information Criteria (AIC) with a maximum 4 lags.

Table 2. ERPT Estimates: IRF Approach

$$\mathbf{x}_t = \sum_{j=1}^{q} \mathbf{B}_j \mathbf{x}_{t-j} + \mathbf{C}\mathbf{u}_t,\ \mathbf{x}_t = [\Delta s_t \quad \Delta y_t \quad \Delta p_t]'$$
$$\phi(j) = E\left(\Delta p_{t+j} | u_{\Delta s,t} = 1, \Omega_{t-1}\right) - E\left(\Delta p_{t+j} | \Omega_{t-1}\right),\ \eta(j) = \sum_{s=0}^{j} \phi(s)$$

	Pre-1990's		Post-1990's	
CPI	*ERPT*	CI	*ERPT*	CI
All item	−0.142	[−0.637, 0.286]	−0.139	[−0.209, −0.081]
Food	−0.125	[−0.400, 0.110]	−0.152	[−0.255, −0.064]
Housing	0.019	[−0.458, 0.508]	−0.064	[−0.145, 0.008]
Apparel	−0.018	[−0.213, 0.172]	0.029	[−0.084, 0.144]
Transportation	0.177	[−0.369, 0.791]	−0.708	[−0.998, −0.475]
Medical Care	0.153	[−0.026, 0.375]	0.013	[−0.138, 0.155]
Energy	0.448	[−0.670, 1.782]	−1.390	[−1.980, −0.893]
All less Energy	−0.036	[−0.345, 0.265]	−0.065	[−0.143, −0.011]
All less Food	−0.007	[−0.469, 0.455]	−0.152	[−0.230, −0.087]
All less F&E	0.028	[−0.259, 0.336]	−0.014	[−0.089, 0.051]

Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the trivariate VAR system with the real exchange rate, the real GDP, and the CPI after log-differencing each variable. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution. The long-run effect *ERPT* is the 20-period ahead response of the level CPI, which is sufficient for the IRFs to get stabilized.

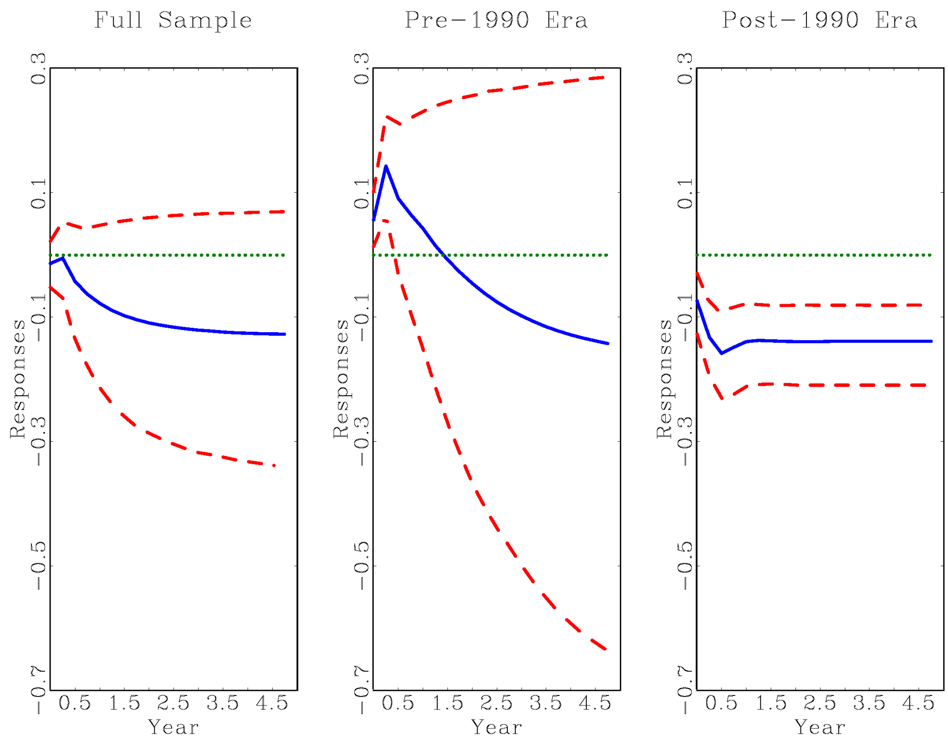
Table 3. ERPT Estimates: VAR Coefficients Approach

$$\mathbf{x}_t = \sum_{j=1}^{q} \mathbf{B}_j \mathbf{x}_{t-j} + \mathbf{C}\mathbf{u}_t,\ \ \mathbf{x}_t = [\Delta s_t \ \ \Delta y_t \ \ \Delta p_t]'$$
$$\Delta p_t = \sum_{j=1}^{q} \beta_{p,j} \Delta p_{t-j} + \sum_{j=0}^{q} \beta_{s,j} \Delta s_{t-j} + \sum_{j=0}^{q} \beta_{y,j} \Delta y_{t-j} + \varepsilon_t$$
$$ERPT = \frac{\sum_{j=0}^{q} \beta_{s,j}}{1-\sum_{j=1}^{q} \beta_{p,j}}$$

	Pre-1990's		Post-1990's	
CPI	*ERPT*	CI	*ERPT*	CI
All item	−0.179	[−1.485, 0.602]	−0.134	[−0.192, −0.081]
Food	−0.137	[−0.326, 0.027]	−0.169	[−0.292, −0.069]
Housing	0.032	[−0.874, 0.861]	−0.063	[−0.150, 0.013]
Apparel	−0.018	[−0.168, 0.141]	0.006	[−0.094, 0.108]
Transportation	0.149	[−0.352, 0.608]	−0.661	[−0.848, −0.479]
Medical Care	0.087	[−0.011, 0.205]	0.016	[−0.145, 0.239]
Energy	0.328	[−0.582, 1.268]	−1.285	[−1.693, −0.882]
All less Energy	−0.031	[−0.379, 0.231]	−0.068	[−0.172, −0.008]
All less Food	−0.018	[−0.896, 0.712]	−0.143	[−0.206, −0.085]
All less F&E	0.026	[−0.377, 0.385]	−0.009	[−0.085, 0.064]

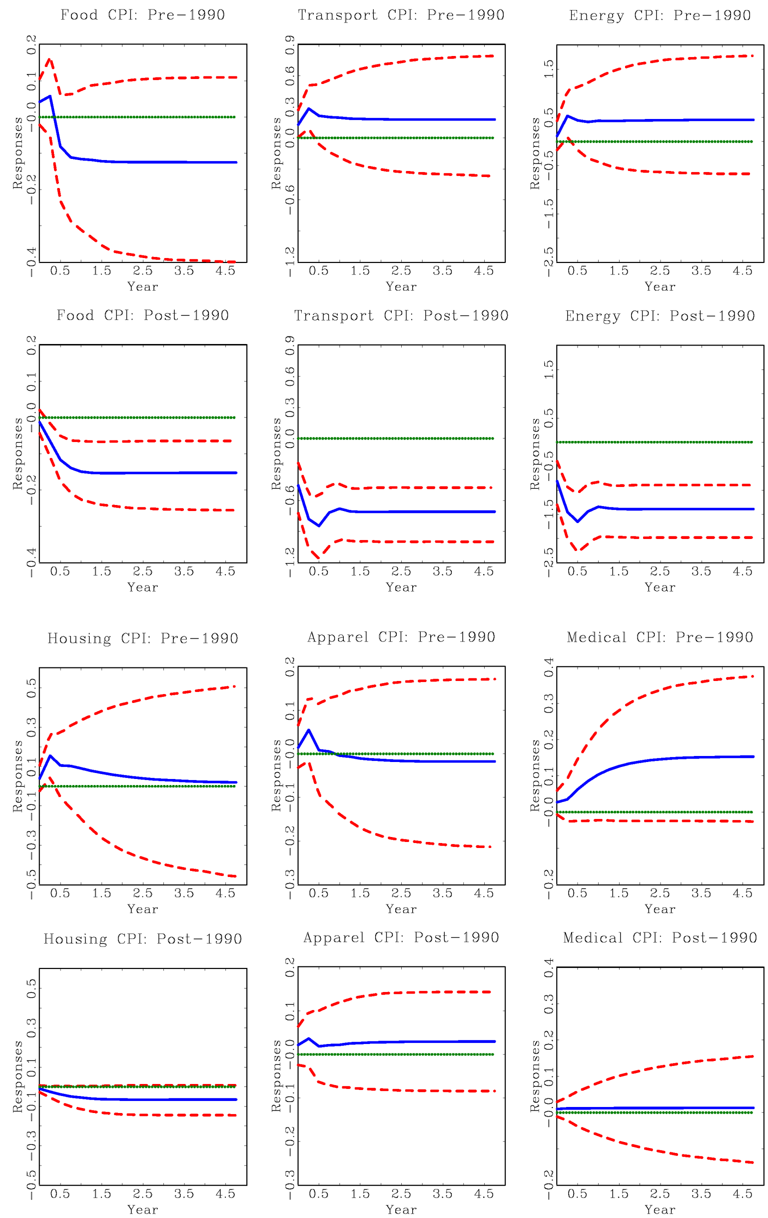
Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the trivariate VAR system with the real exchange rate, the real GDP, and the CPI after log-differencing each variable. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. *ERPT* is constructed using the structural coefficient estimates from the VAR system. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution.

Figure 1. CPI Responses to the Exchange Rate Shock



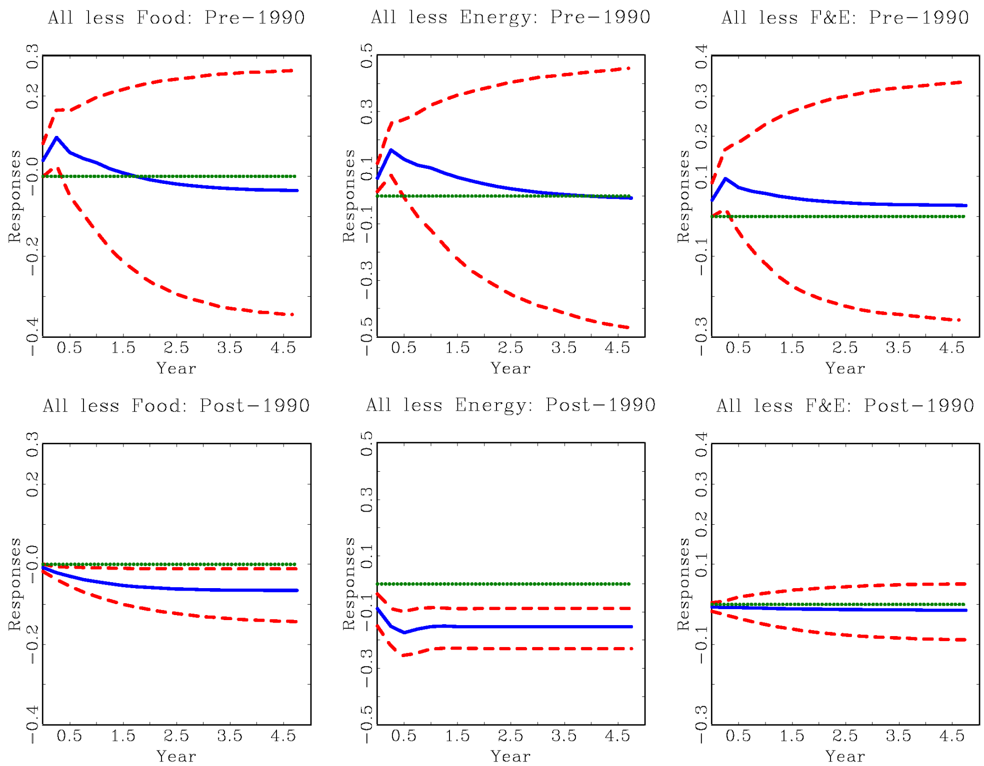
Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the trivariate VAR system with the real exchange rate, the real GDP, and the CPI after log-differencing each variable. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution.

Figure 2. Disaggregated CPI Responses to the Exchange Rate Shock



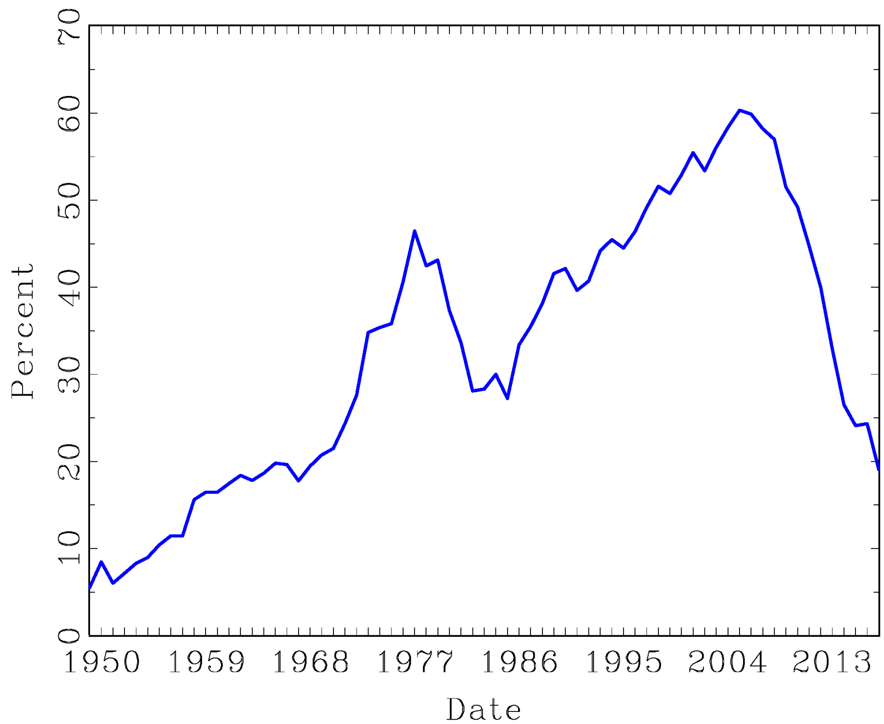
Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the trivariate VAR system with the real exchange rate, the real GDP, and the CPI after log-differencing each variable. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution.

Figure 3. Special CPI Responses to the Exchange Rate Shock



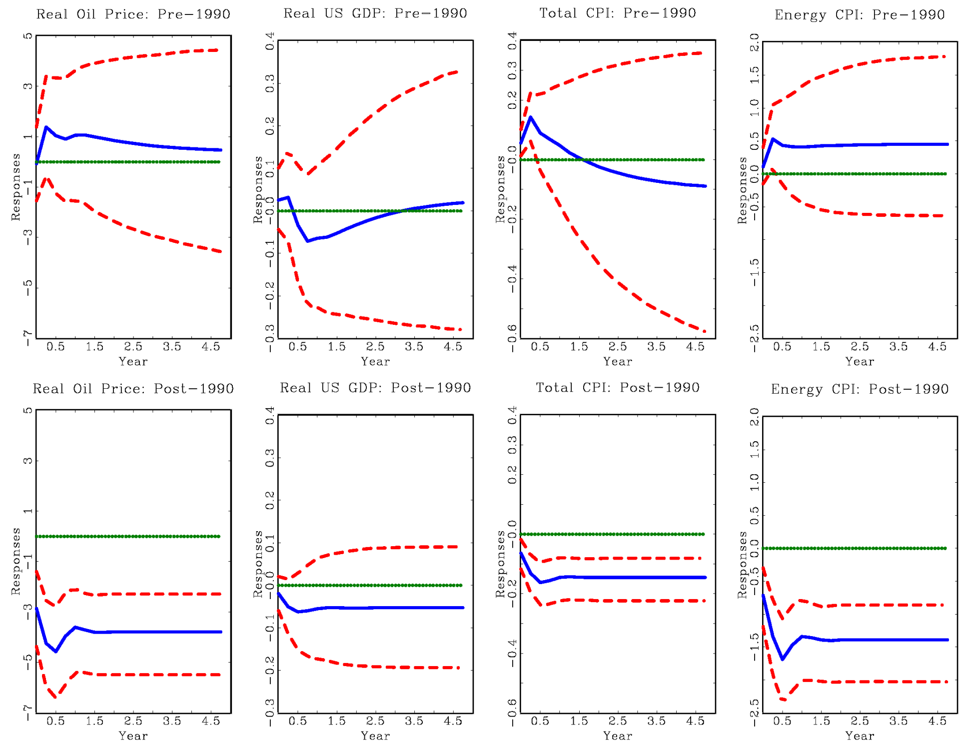
Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the trivariate VAR system with the real exchange rate, the real GDP, and the CPI after log-differencing each variable. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution.

Figure 4. Consumption Dependency on Foreign Petroleum



Note: We report the share of the U.S. petroleum net imports. We obtained the data from the Energy Information Agency.

Figure 5. IRF Analysis with the Real Oil Price



Note: We estimate the impulse-response function of the level CPI to the real exchange rate shock from the quadvariate VAR system with the real exchange rate, the real oil price, the real GDP, and the CPI. We log-differenced all variables prior to estimations. The pre-1990 sample period is from 1973:II to 1989:IV, while the post-1990 begins in 1990:I and ends in 2017:IV. We set the number of lag to 2. The 90% confidence bands were constructed by taking the 5 and 95 percentiles from 5,000 nonparametric bootstrap simulations using the empirical distribution.