Journal of International Economics 81 (2010) 48–60

Contents lists available at ScienceDirect

Journal of International Economics

journal homepage: www.elsevier.com/locate/jie

Accounting for persistence and volatility of good-level real exchange rates: The role of sticky information

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ARTICLE INFO

Article history:
Received 28 May 2008
Received in revised form 18 December 2009
Accepted 14 January 2010

JEL classification:
E31
F31
D40

Keywords:
Good-level real exchange rates
Law of one price
Sticky information
Dynamic panel

ABSTRACT

Volatile and persistent real exchange rates are observed not only in aggregate series but also in micro-price data at the retail level. Kehoe and Midrigan (2007) recently showed that, under a standard assumption on nominal price stickiness, empirical frequencies of micro price adjustment cannot replicate the time series properties of the Law of One Price (LOP) deviations. We extend their sticky price model by combining good-specific price adjustment with information stickiness in the sense of Mankiw and Reis (2002). Our framework allows for multiple cities within a country. Using a panel of U.S.–Canadian city pairs, we estimate a dynamic price adjustment process for 165 individual goods. Under a reasonable assumption on the money growth process, we show that the model matches the persistence of the LOP deviation for the median good and accounts for the majority of its volatility when information updates occur every 12 months.

1. Introduction

Aggregate real exchange rates exhibit persistence and volatility much higher than what economists believe is consistent with a plausible degree of price rigidity. The time-dependent pricing model, under local currency pricing, offers a convenient theoretical link between price stickiness and the stochastic properties of real exchange rates. Chari et al. (2002, CKM) show that to generate the observed persistence of CPI-based aggregate real exchange rates, prices need to be exogenously fixed for at least one year. This degree of price stickiness, however, appears implausible based on recent evidence by Bils and Klenow (2004) who find a median duration between price changes of only 4.3 months in U.S. micro-data.

Using a broader sample of the Economist Intelligence Unit (EIU) micro-price data than employed here, Crucini and Shintani (2008) find the half-life of deviations from the Law of One Price (LOP) for the median good in the neighborhood of 18 months, considerably lower than the consensus 3–5 year half-lives of aggregate real exchange rates. An important theoretical contribution along similar lines is Kehoe and Midrigan (2007) who allow different price stickiness across individual goods and show that the persistence in LOP deviations is equal to 'the Calvo parameter,' the probability of price non-adjustment at the good level. Their empirical analysis using real exchange rates of 66 individual goods across the U.S. and four European countries shows that the frequency of no price adjustment is higher for goods that exhibit more persistent deviations from the LOP, as suggested by the theoretical model. However, the persistence puzzle is still not resolved in the sense that the observed frequencies of price changes are too high to replicate the persistence of real exchange rates for most goods in the cross-section. In addition, the model does not match the time series variability of LOP deviations observed in the micro-data. Their results point to the need to break the tight link between the frequency of price adjustment and the persistence of LOP deviations predicted by the standard Calvo-type sticky price model.

Our analysis differs from Kehoe and Midrigan (2007) in several ways. First, we break the tight link between the Calvo parameter and LOP persistence by extending the Kehoe-Midrigan model to allow for persistent money growth and information stickiness. Information stickiness is the assumption that only a fraction of firms update their information set each period. Here, LOP persistence arises from the convolution of price adjustment timing and information updating. In the macroeconomic literature, Mankiw and Reis (2002) show that a model of information stickiness, or inattentiveness, is capable of

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0022-1996/$ – see front matter
doi:10.1016/j.jinteco.2010.01.003

explaining the observed slow response of aggregate inflation to monetary shocks much better than sticky prices alone. When the information stickiness augments the Calvo-type sticky price mechanism, less frequent information updating leads to higher price persistence, at a given frequency of price adjustment (Dupor et al., forthcoming, DKT). With plausible assumptions on international money growth processes, a similar effect takes place to increase both the persistence and volatility of real exchange rates.[1]

Second, our theoretical model allows for multiple cities in each country and for long-run price deviations between the cross-border city pairs to differ by good and city pair. As such, our model allows us to exploit an international retail price survey at the city level which records local currency prices of individual goods and services spanning most of the CPI basket. Because this survey is conducted by a single agency, the EIU, we expect more comparability of the products among international cities than is true of national CPI surveys, bringing the data more in line with the spirit of the model. Using this survey, from 1990 to 2005, we expand the number of products from 66 used in Kehoe and Midrigan (2007) to 165. We also increase the number of locations from five countries (Austria, Belgium, France, Spain and the U.S.) to 52 U.S.–Canadian city pairs.

Third, we examine the effect of the exclusion of sales on the performance of the model. Recently, Nakamura and Steinsson (2008) claim that the evidence of fast price adjustment reported by Bils and Klenow (2004) may be strongly influenced by the presence of sales, or other temporary price reductions. Focusing on regular price changes, Nakamura and Steinsson (2008) find that the median duration between price changes increases to the range of 8 to 11 months. Since prices are stickier based on this alternative definition of price change, it elevates the Calvo model's ability to account for important features of the data. [2] This improvement is subject to the caveat that we do not explicitly model sales.

The shortcomings of the Calvo model highlighted in Kehoe and Midrigan (2007) persist in the context of the EIU data with or without the adjustment of the frequency of price changes for temporary sales. In contrast, the Calvo model extended to allow for information stickiness fully accounts for LOP persistence of the median good and the majority of its volatility when the average duration between information updates is about one year, when we calibrate to the observed frequency of regular price changes.[3]

2. The model

Trade is over a continuum of goods across multiple cities located in two countries. Under monopolistic competition, firms set prices in local currency to satisfy demand for a particular good in a particular city. A representative agent in each country chooses consumption over an infinite horizon subject to a cash-in-advance (CIA) constraint. In what follows, the U.S. and Canada represent the home and foreign country. Due to the symmetry of the model, we mostly focus on the equations of the United States.

The lowest level of aggregation is the brand, z of a particular good. U.S. brands of each good are indexed $z\in[0, 1/2]$, while those of Canada are indexed $z\in(1/2, 1]$. Integrating over brands, we have the CES index for consumption of good j, given by

$$c_t(j,l) = \left[\int_0^1 c_t(j,l,z)^{\frac{\theta-1}{\theta}}dz\right]^{\frac{\theta}{\theta-1}}, \tag{1}$$

where $c_t(j, l, z)$ is consumption of a brand z of good j in U.S. city l. CES aggregation across U.S. cities $l\in[0, 1]$, gives U.S. national consumption of good j

$$c_t(j) = \left[\int c_t(j,l)^{\frac{\theta-1}{\theta}}dl\right]^{\frac{\theta}{\theta-1}}, \tag{2}$$

and further CES aggregation across goods gives aggregate U.S. consumption, c_t,

$$c_t = \left[\int c_t(j)^{\frac{\theta-1}{\theta}}dj\right]^{\frac{\theta}{\theta-1}}. \tag{3}$$

2.1. Households

As in Kehoe and Midrigan (2007), markets for state-contingent money claims are complete. The asset structure is represented with one period state-contingent bonds. Let $B(s^t, s_{t+1})$ denote the number of units of bonds in U.S. dollars. These bonds are purchased in period t, indexed by the history of events up to period t, s^t, and pay one U.S. dollar per unit purchased, if the state s_{t+1} is realized in period $t+1$. We suppress the state and denote these holdings as B_{t+1} for U.S. households and B^*_{t+1} for Canadian households with associated nominal stochastic discount factor across adjacent time periods denoted, $\Upsilon_{t,t+1}$.[4] Also, $\Upsilon_{t,t+h}$ is used by firms, regardless of their country of origin, to discount profits earned in period $t+h$ back to the period t.

Households maximize the discounted sum of utility subject to an intertemporal budget constraint and a CIA constraint. The maximization problem for U.S. households is

$$\max\ \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t(\ln c_t - \chi n_t), \tag{4}$$

$$\text{s.t. } M_t + \mathbb{E}_t(\Upsilon_{t,t+1}B_{t+1}) = R_{t-1}W_{t-1}n_{t-1} + B_t + (M_{t-1} - P_{t-1}c_{t-1}) + T_t + \Pi_t, \tag{5}$$

$$M_t \ge P_t c_t, \tag{6}$$

where n_t is hours of work, and $\mathbb{E}_t(\cdot)$ denotes the expectation operator conditional on the information available in period t. We assume that $\chi>0$ and β is between 0 and 1. The left-hand-side of the intertemporal budget constraint (5) represents the U.S. dollar value of household wealth brought into the beginning of period $t+1$. It consists of cash holding, M_t, and bond holdings, B_{t+1}. As shown on the right-hand-side of Eq. (5), the household receives nominal labor income $W_{t-1}n_{t-1}$ in period $t-1$ which earns gross nominal interest R_{t-1} until period t.[5] The household carries nominal bonds in amount B_t and cash holding remaining after consumption expenditures $(M_{t-1}-P_{t-1}c_{t-1})$ into period t. Finally, T_t and Π_t are nominal lump sum transfers from the U.S. government and nominal profits of firms operating in the U.S., respectively.

Eq. (6) is the CIA constraint. The aggregate price P_t is given by $P_t = \left[\int P_t(j)^{1-\theta}dj\right]^{\frac{1}{1-\theta}}$, where $P_t(j)$ is the aggregate price index for good j; it is a CES aggregate over city-specific prices for that good: $P_t(j) = \left[\int P_t(j,l)^{1-\theta}dl\right]^{\frac{1}{1-\theta}}$. The price index for good j in a particular city l used in this aggregation is given by

$$P_t(j,l) = \left[\int P_t(j,l,z)^{1-\theta}dz\right]^{\frac{1}{1-\theta}}. \tag{7}$$

[1] In a related study, Bacchetta and van Wincoop (2006) also use a partial information structure to model nominal exchange rate determination.

[2] Our working paper includes results using the Bils and Klenow frequencies, which are omitted here due to space considerations.

[3] When calibrated to the frequency of price changes including sales the average duration between information updates needed to match LOP persistence rises to between 14 and 20 months.

[4] As Kehoe and Midrigan (2007) argue, the choice of currency denomination of bonds is irrelevant when the markets for state-contingent money claims are complete.

[5] We assume that the government pays interest rate $R_t(=1/\mathbb{E}_t\Upsilon_{t,t+1})$ on labor income in period t. This assumption allows households' intratemporal first-order condition to be undistorted.