Deviations from the Law of One Price Across Cities: Testing for a Border Effect in Persistence and Volatility

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Abstract

In this paper, we test for border effects in deviations from the law of one price. With respect to the AR(p) representation of deviations from the law of one price, we test for a border effect in both the sum of the autoregressive coefficients (persistence) and the variance of the disturbance term (volatility). Using consumer price index data for a large number of categories of consumer goods and services for Canadian and U.S. cities from Engel and Rogers (1996), we first use the Hansen (1999) grid-bootstrap procedure to generate median-unbiased estimates of the persistence and volatility in deviations from the law of one price for every city pair and CPI category. We then estimate cross-section regression models across city pairs for the estimates of persistence and volatility for each CPI category, and we find that (after controlling for the distance between cities) crossing the Canadian-U.S. border leads to significant increases in both persistence and volatility for nearly all of the CPI categories. We also document a border effect in both persistence and volatility using the aggregate consumer price index data for European cities from Engel and Rogers (2001).

JEL classifications: C22; D40; F31; F41

Key words: Autoregressive model; Grid bootstrap; Law of one price; Persistence; Volatility

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1. Introduction

Partly stimulated by the earlier findings of, among others, Isard (1977), Richardson (1978), Dornbusch (1987), and Giovannini (1988), the last decade has witnessed the emergence of a substantial empirical literature investigating international deviations from the law of one price (LOOP) and their implications for models of real exchange rate behavior. A popular empirical approach follows Engel (1993) and measures the volatility in deviations from the LOOP. Interestingly, this approach documents substantial volatility in deviations from the LOOP across countries. For example, Engel (1993) shows that the volatility in deviations from the LOOP for similar goods across countries—even putatively traded goods—is much greater than the volatility in relative prices of different goods within a country. This result suggests that fluctuations in real exchange rates (deviations from purchasing power parity) are not primarily due to changes in the relative prices of non-traded goods within countries and thus is at odds with the "standard" explanation of real exchange rate fluctuations offered by Balassa (1964) and Samuelson (1964). Other papers measuring the volatility in deviations from the LOOP include the well-known study by Engel and Rogers (1996), who use CPI data for 14 categories of consumer goods and services for U.S. and Canadian cities; Parsley and Wei (2001a), who use the prices of 27 traded goods for U.S. and Japanese cities; Engel and Rogers (2001), who use aggregate and disaggregated CPI data for European cities and countries; Parsley and Wei (2001b), who use the prices of 95 very disaggregated goods for 83 cities around the world. In line with Engel (1993), these studies document a large "border effect": after controlling for distance, the volatility in deviations from the LOOP—even for goods that are typically classified as traded—is much higher between locations that lie across a border.

While the volatility in deviations from the LOOP is found to increase when crossing a border, whether or not the persistence in these deviations from the LOOP increases when a crossing a border is an unanswered question. The purpose of this paper is to shed light on this issue. As emphasized by Engel (2000a,b), the volatility in deviation from the LOOP has important implications for the theoretical models in international macroeconomics, especially those of the so-called "new open-economy macroeconomics" (NOEM) variety, which rely on an optimization-based dynamic general equilibrium framework combined with some form of sticky prices and/or wages in some sectors of the economy.[1] The persistence in deviations from the LOOP also has important implications for these models. More specifically, measuring the persistence in deviations from the LOOP is important for determining whether the assumption of producer-currency pricing (PCP), as used in the seminal NOEM model of Obstfeld and Rogoff (1995), or the assumption of local currency pricing (LCP), as used more recently by Betts and Devereaux (1996, 2000) and Bachetta and van Wincoop (2000), is more appropriate in NOEM models. Under PCP, producers set prices in the domestic currency, so that the law of one price prevails at all times (and hence, there is no allowance for deviations from the LOOP). This opens up an "expenditure-switching" channel whereby a domestic monetary expansion causes a depreciation of the domestic currency that increases the demand for domestically produced goods. Such an expenditure-switching channel also characterizes the traditional Mundell-Fleming-Dornbusch model (Engel, 2000), and it provides a theoretical argument against fixing the nominal exchange rate (Obstfeld and Rogoff, 2002; Engel, 2002a,b). Under LCP, produces set price in the local currency, the LOOP no longer holds, allowance is made for deviations from the LOOP, and the expenditure-switching channel is eliminated.[2] This has implications for monetary policy, as LCP can imply that a degree of nominal exchange rate stability is optimal (Engel, 2002b).

Given the important theoretical implications of deviations from the LOOP, the present paper aims to develop a fuller understanding of the nature of the empirical deviations from the LOOP. While the extant literature tests for a border effect in volatility, we extend the extant literature by testing for a border effect with respect to both the persistence and volatility in deviations from the LOOP. As a scalar measure of persistence, we focus on the sum of the autoregressive coefficients (SARC) in the AR(p) representation of deviations from the LOOP. Andrews and Chen (1994) argue that the SARC is an informative metric for the persistence in an AR(p) process, as it measures the cumulative impulse

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[1] See Lane (2001), Lane and Ganelli (2002), and Bowman and Doyle (2003) for useful surveys of the NOEM literature.

[2] Under local-currency pricing, there is zero exchange rate pass-through.

response function (the sum of the impulse response function over all time horizons).[3] It is well known that OLS estimates of the SARC are biased, so we use the Hansen (1999) grid-bootstrap procedure to generate median-unbiased estimates of the SARC. We measure volatility as the variance of the disturbance term of the fitted AR(*p*) process. We also use the Hansen (1999) grid-bootstrap procedure to calculate the lower bound of the 95% confidence interval for the SARC, as this represents the lowest degree of persistence in deviations from the LOOP consistent with the data.

We measure the persistence and volatility in deviations from the LOOP using two data sets. The first is the monthly data from Engel and Rogers (1996) that includes disaggregated CPI data for 14 categories of consumer goods and services and a number of Canadian and U.S. cities. For each category of goods and services, we use the Hansen (1999) grid-bootstrap procedure to calculate a median- unbiased estimate of the SARC for every city pair. Analogous to Engel and Rogers (1996), we then test for a border effect in persistence by regressing the median-unbiased estimates of the SARC for each city pair on the distance between cities and a border dummy. For 11 of the 14 categories of goods and services, we find that (after controlling for the distance between cities) crossing the Canadian-U.S. border significantly increases the SARC and thus the persistence in deviations from the LOOP. We obtain similar results when we measure persistence using the lower bound of the 95% confidence interval for the SARC. For each CPI category, we also calculate the volatility in deviations from the LOOP for every city pair, and like Engel and Rogers (1996), we find that crossing the Canadian-U.S. border significantly increases the volatility in deviations from the LOOP. The second data set is the monthly aggregate CPI data for a large number of European cities from Engel and Rogers (2001). We again document a significant border effect in both the persistence and volatility in deviations from the LOOP.

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[3] Andrews and Chen (1994) also point out that the sum of the AR coefficients is directly related to the spectrum at frequency zero. They regard the SARC as a more informative measure of persistence than the largest root of the AR(*p*) process, as two AR(*p*) processes with identical largest roots can nonetheless have very different persistence properties. Rogers and Jenkins (1995) measure persistence in deviations from the LOOP using country-wide consumer price data for 54 categories of goods and services in the U.S. and Canada using the largest AR root and the procedure in Stock (1991) to construct confidence intervals for the largest AR root.

The rest of the paper is organized as follows: Section 2 outlines the econometric methodology, Section 3 reports our empirical results, and Section 4 concludes.

2. Econometric Methodology

Consider the following AR(*p*) process for $y_t^{i,j}$:

$$y_t^{i,j} = \mu^{i,j} + \alpha_1^{i,j} y_{t-1}^{i,j} + \alpha_2^{i,j} y_{t-2}^{i,j} + \ldots + \alpha_p^{i,j} y_{t-p}^{i,j} + e_t^{i,j} , \qquad (1)$$

where $t = 1,2,\ldots,T$, $y_t^{i,j}$ is the deviation from the LOOP between cities i and j ($i \neq j$), and $e_t^{i,j}$ is distributed independently and identically with mean zero and variance $(\sigma_e^{i,j})^2$. In Section 3 below, $y_t^{i,j}$ is the natural logarithm of the ratio of the CPI for a given component in city i to the CPI for a given component in city j . As mentioned in the introduction, Andrews and Chen (1994) argue that the sum of the AR coefficients (SARC), $\alpha^{i,j} = \sum_{k=1}^{p} \alpha_k^{i,j}$, is an informative scalar measure of the persistence in the AR process. Using $\alpha^{i,j} = \sum_{k=1}^{p} \alpha_k^{i,j}$, equation (1) can be equivalently expressed in terms of the familiar augmented Dickey and Fuller (1981, ADF) and Said and Dickey (1984) regression model,

$$y_t^{i,j} = \tilde{\mu}^{i,j} + \alpha^{i,j} y_{t-1}^{i,j} + \sum_{k=1}^{p-1} \beta_k^{i,j} \Delta y_{t-k}^{i,j} + e_t^{i,j} , \qquad (2)$$

where $\Delta y_t^{i,j} = y_t^{i,j} - y_{t-1}^{i,j}$. We use $\alpha^{i,j}$ in equation (2) to measure the persistence in deviations from the LOOP between two cities, and we measure the volatility in deviations from the LOOP using $(\sigma_e^{i,j})^2$. In our applications in Section 3 below, we use the modified AIC criterion of Ng and Perron (2001) to select the lag (p) in equation (2).[4]

There are well-known econometric problems associated with the estimation of $\alpha^{i,j}$ in equation (1) or (2). OLS point estimates of $\alpha^{i,j}$ suffer from potentially large biases, especially when $\alpha^{i,j}$ is near

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[4] Given that we are working with CPIs, we cannot measure absolute deviations from the LOOP, but only relative deviations. This provides a rationale for including a constant term in equations (1) and (2). A constant term also allows for a permanent wedge between prices due, for example, to various types of transport costs, tax differentials, etc.

unity (Hurwicz, 1950; Kendall, 1954; Shaman and Stine, 1988). In addition, conventional bootstrap procedures fail to generate confidence intervals for $\alpha^{i,j}$ with correct first-order asymptotic coverage (Basawa et al., 1991; Hansen, 1999). In light of these problems, we rely on the recently developed grid-bootstrap procedure of Hansen (1999) to compute median-unbiased point estimates of $\alpha^{i,j}$ and lower bounds of 95% confidence interval for $\alpha^{i,j}$ in equation (2) that are valid even when $\alpha^{i,j}$ is near unity. As mentioned in the introduction, the lower bound of the 95% confidence interval for $\alpha^{i,j}$ represents the lowest degree of persistence in deviations from the LOOP consistent with the data. We use the Hansen (1999) procedure to compute median-unbiased estimates of $\alpha^{i,j}$ instead of the method suggested by Andrews and Chen (1994), as the Andrews and Chen (1994) procedure assumes $\alpha^{i,j} \leq 1$, while the Hansen (1999) method does not impose this restriction.

In order to calculate a 95% confidence interval for $\alpha^{i,j}$ in equation (2), the Hansen (1999) grid-bootstrap procedure works as follows. Consider a grid of values for $\alpha^{i,j}$, $\alpha_b^{i,j}$ $(b = 1,\ldots,B)$, covering $\hat{\alpha}^{i,j}$, where $\hat{\alpha}^{i,j}$ is the OLS estimator of $\alpha^{i,j}$ in equation (2). (The grid contains 200 points in our applications in Section 3 below.) In order to estimate the data-generating process for each $\alpha_b^{i,j}$, we estimate equation (2) with $\alpha^{i,j}$ restricted to $\alpha_b^{i,j}$ using restricted OLS for each $\alpha_b^{i,j}$. The restricted OLS parameter estimates, together with re-sampled restricted OLS residuals, are then used to build up a large number of pseudo-samples for each $\alpha_b^{i,j}$. (We use 999 in our applications in Section 3 below.) For each of the 999 pseudo-samples for each $\alpha_b^{i,j}$, we calculate the *t*-statistic, $t_b^* = (\hat{\alpha}_b^{i,j*} - \alpha_b^{i,j}) / s(\hat{\alpha}_b^{i,j*})$, where $\hat{\alpha}_b^{i,j*}$ is the OLS estimate of $\alpha^{i,j}$ for the pseudo-sample and $s(\hat{\alpha}_b^{i,j*})$ is its standard error. We then sort the *t*-statistics, giving us an empirical distribution of *t*-statistics for each $\alpha_b^{i,j}$, from which we can calculate the 0.025 and 0.975 quantiles of *t*-statistics for each $\alpha_b^{i,j}$. The upper bound for the 95% confidence interval for $\alpha^{i,j}$ is the $\alpha_b^{i,j}$ grid value such that $(\hat{\alpha}^{i,j} - \alpha_b^{i,j}) / s(\hat{\alpha}^{i,j}) = t_{b,0.025}^*$, while the lower bound is the $\alpha_b^{i,j}$ grid value such that $(\hat{\alpha}^{i,j} - \alpha_b^{i,j}) / s(\hat{\alpha}^{i,j}) = t_{b,0.975}^*$. In Monte Carlo simulations, Hansen (1999) finds that the grid-bootstrap procedure generates confidence intervals with good coverage in finite samples. It is straightforward to calculate a median-unbiased estimator of $\alpha^{i,j}$ ($\hat{\alpha}_{MU}^{i,j}$) using the procedure described above, with the exception that we use the 0.50 quantile in place of the 0.025 and 0.975 quantiles. An estimate of $(\sigma_e^{i,j})^2$ $[(\hat{\sigma}_e^{i,j})^2]$ is readily obtained using the fitted residuals in equation (2). We use the Hansen (1999) grid-bootstrap procedure to calculate median-unbiased estimates of $\alpha^{i,j}$ ($\hat{\alpha}_{MU}^{i,j}$) and lower bounds of the 95% confidence intervals for $\alpha^{i,j}$ ($\hat{\alpha}_{LB}^{i,j}$).

In order to formally test for a border effect in persistence based on the median-unbiased estimates of $\alpha^{i,j}$, we estimate the following cross-section regression model that is similar to the cross-section regression models for volatility used by Engel and Rogers (1996):

$$\hat{\alpha}_{MU}^{i,j} = \sum_{m=1}^{M} \gamma_m D_m + \delta x_{i,j} + \phi B_{i,j} + u_{i,j}, \tag{3}$$

where $x_{i,j}$ is the distance between cities i and j,[5] $B_{i,j}$ is a border dummy variable that takes on the value of one when cities i and j are in different countries, and D_m is a dummy variable for each city (M is the total number of cities) that takes on a value of one for cities i and j. It is important to control for distance when testing for a border effect, as transport costs are likely to be higher for cities that are farther apart. We follow Engel and Rogers (1996) and include a dummy variable for each city in equation (3). Engel and Rogers (1996) argue that it is appropriate to include a dummy variable for each city in the regression model due to, for example, measurement error or seasonalities in some cities. We are primarily interested in the significance of ϕ in equation (3), as it allows us to test for a border effect in persistence after controlling for the distance between cities. We also estimate equation (3) using $\hat{\alpha}_{LB}^{i,j}$ in place of $\hat{\alpha}_{MU}^{i,j}$, and we test for border effect in volatility by using $(\hat{\sigma}_e^{i,j})^2$ in place of $\hat{\alpha}_{MU}^{i,j}$ in equation (3). For all of the cross-section regressions, we use the White (1980) procedure to compute heteroskedasticity-consistent standard errors and *t*-statistics.

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[5] As in Engel and Rogers (1996), distance is measured using the natural logarithm of the great-circle distance.

Engel and Rogers (1996) focus on the volatility of deviations from the LOOP. Their primary measure of volatility is the standard deviation of $y_t^{i,j} - y_{t-1}^{i,j}$.[6] This is equivalent to measuring volatility as $\sigma_e^{i,j}$ in equation (2) after imposing the restrictions, $\alpha^{i,j} = 1$ and $\beta_k^{i,j} = 0$ ($k = 1,\ldots, p-1$), so that $y_t^{i,j}$ is modeled as a random walk. We use a more general approach in order to consider potential differences in both persistence ($\alpha^{i,j}$) and volatility [$(\sigma_e^{i,j})^2$] across pairs of cities.[7]

3. Empirical Results

3.1. Engel and Rogers (1996) Canadian and U.S. City Data

Engel and Rogers (1996) employ monthly CPI data disaggregated into 14 categories of goods and services for 9 Canadian and 14 U.S. cities covering the period June 1978 to December 1994. They find that the volatility of deviations from the LOOP between U.S. city pairs is generally slightly higher than between Canadian city pairs. Most dramatically, deviations from the LOOP for cross-border city pairs have much higher volatility than deviations between cities within the same country (the border effect). We use the Engel and Rogers (1996) data for nine Canadian cities (Calgary, Edmonton, Montreal, Ottawa, Quebec, Regina, Toronto, Vancouver, Winnipeg) and the four U.S. cities for which continuous monthly data are available (New York, Chicago, Philadelphia, Los Angeles). Using this data, we measure the persistence and volatility in deviations from the LOOP for each city pair for the All items CPI and each CPI category using the methodology described in Section 2 above, where the deviations are measured as the log-level of the ratio of the U.S. dollar price of a given consumption category in two different cities. We then formally test for a border effect with respect to both persistence and volatility for the All items CPI and each CPI category using the cross-section regression model, equation (3).

Figure 1 provides a visual perspective on the differences in persistence between city pairs within Canada or the U.S. and pairs of cities lying across the border. The solid line in each panel portrays the

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[6] Engel and Rogers actually use $y_t^{i,j} - y_{t-2}^{i,j}$, but this difference is not crucial.

[7] Engel and Rogers (1996) note that they obtain similar results when they compute volatility using the in-sample forecast error from an AR(6) model fitted to $y_t^{i,j}$, and this approach is closer to ours.

empirical cumulative density function (cdf) of $\hat{\alpha}_{MU}^{i,j}$ values for Canadian-Canadian and U.S.-U.S. city pairs, while the dashed line graphs the empirical cdf for Canadian-U.S. city pairs.[8] With the exception of Shelter, the cdf corresponding to city pairs within Canada or the U.S. is well to the left of the cdf for city pairs lying across the border for the All items CPI and each CPI category. This indicates that we are more likely to find smaller $\hat{\alpha}_{MU}^{i,j}$ values for city pairs within Canada or the U.S. and larger $\hat{\alpha}_{MU}^{i,j}$ values for city pairs in different countries. Figure 2 shows that a similar pattern holds for the empirical distribution of $(\hat{\sigma}_{e}^{i,j})^2$ values: for the All items CPI and each CPI category, the cdf for city pairs within Canada or the U.S. is well to the left of the cdf for city pairs in different countries. Together, Figures 1 and 2 suggest that greater persistence and volatility is associated with crossing the Canadian-U.S. border. We now test this proposition more formally using the cross-section regression model, equation (3).

The cross-section regression results are reported in Table 1. From column (3) of Table 1, we see evidence of a significant border effect with respect to $\hat{\alpha}_{MU}^{i,j}$ for the All items CPI and 11 of the 14 CPI categories, as the border coefficient is significant at conventional levels in equation (3) in these cases. The border coefficients range from 0.002-0.036. The distance coefficient is significant at conventional levels for only 4 of the 14 categories. It is interesting to note that the border coefficients are significant for a number of categories, such as Food at home, that are primarily composed of putatively traded goods. To gain some additional insight into the quantitative importance of the border effect, we consider the half-life, the number of months required for a unit shock to the deviations from the LOOP to decrease by 0.5. For an AR(p) process, this can be approximated by $\ln(0.5)/\ln(\alpha^{i,j})$. Suppose we begin with the mean $\hat{\alpha}_{MU}^{i,j}$ value for U.S.-U.S. city pairs and the All items CPI of 0.956. The corresponding half-life for $\hat{\alpha}_{MU}^{i,j} = 0.956$ is 15.404 months. According to the estimated border coefficient of 0.031 in Table 1, crossing the border will increase $\alpha^{i,j}$ to 0.987, with a corresponding half-life of 52.792 months. This

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[8] The empirical cumulative density functions were calculated using the nonparametric density estimation procedure in Hansen (2004).

represents an increase in the half-life by a factor of 3.439. This indicates that the border effect with respect to $\alpha^{i,j}$ can be economically, as well as statistically, significant in Table 1.

From column (6) of Table 1, we see that there is strong evidence of a border effect with respect to $(\hat{\sigma}_{e,MU}^{i,j})^2$, as the border coefficient is significant at the 1% level for the overall CPI and all of the CPI categories. This matches the Engel and Rogers (1996) result. Overall, we find strong evidence of a border effect with respect to both persistence and volatility for the All items CPI and nearly all of the CPI categories in Table 1.

Table 2 reports cross-section regression results when $\hat{\alpha}_{LB}^{i,j}$ is used in place of $\hat{\alpha}_{MU}^{i,j}$ in equation (3). We see from column (3) of Table 2 that there is significant evidence of a border effect with respect to $\hat{\alpha}_{LB}^{i,j}$ for the All items CPI and 13 of the 14 CPI categories. The border coefficients range from 0.012-0.080. Note that the estimated coefficients in column (3) of Table 2 are larger than the corresponding entries in column (3) of Table 1, so that the measured border effect is even stronger when we use $\hat{\alpha}_{LB}^{i,j}$ in place of $\hat{\alpha}_{MU}^{i,j}$. The distance coefficient is significant at conventional levels for only 4 of the 14 categories.

3.2. Engel and Rogers (2001) European City Data

Engel and Rogers (2001) investigate deviations from the LOOP based on overall CPI data for 55 locations across Europe. The city-level data employed by Engel and Rogers (2001) are only available for the overall CPI, and the data are monthly and span the period from March 1981 to July 1997. The econometric methodology used by Engel and Rogers (2001) is similar to that of Engel and Rogers (1996), and Engel and Rogers (2001) also document a border effect.[9] We employ the Engel and Rogers (2001) data to test for a border effect with respect to both the persistence and volatility in deviations from purchasing power parity across European cities. There are four countries (Germany, Italy, Spain, Switzerland) for which overall CPI data are available for different cities within the country. For seven

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[9] Due to the nature of the data used in Engel and Rogers (2001), they can separate out the border effect into the part due to changes in nominal exchange rates and that due to a “real” border effect. They find the former to be the most important.

other countries (France, Austria, Belgium, Denmark, Luxembourg, Netherlands, Portugal), CPI data for the country as a whole are only available, and we follow Engel and Rogers (2001) and associate the overall CPI with the capital city of the country. Again following Engel and Rogers (2001), we measure deviations from the LOOP between two given locations as the difference in the log-levels of the CPIs expressed in U.S. dollars.

Table 3 reports the cross-section regression results. From column (2) of Table 3, we see that crossing a border brings about a significant increase in persistence. This hold for both $\hat{\alpha}_{MU}^{i,j}$ and $\hat{\alpha}_{LB}^{i,j}$. Note that the border coefficients for $\hat{\alpha}_{MU}^{i,j}$ and $\hat{\alpha}_{LB}^{i,j}$ in Table 3 are smaller than the corresponding coefficient estimates for the All items CPI reported in Tables 1 and 2. The border effect with respect to persistence thus appears smaller for European borders than for the Canadian-U.S. border. From column (5) of Table 3, we see that crossing a European border results in a significant increase in volatility. Observe that the border coefficient for $(\hat{\sigma}_{e}^{i,j})^2$ in Table 3 is larger than the corresponding coefficient estimate for the All items CPI reported in Table 1. The border effect with respect to volatility thus appears larger for European border than for the Canadian-U.S. border.

4. Conclusion

The recent empirical literature demonstrates a border effect with respect to the volatility of deviations from the LOOP. We add value to the recent empirical literature by providing evidence of a border effect with respect to persistence, as well as volatility, in deviation from the LOOP. With respect to theoretical models in international macroeconomics, the highly persistent and highly volatile nature of deviations from the LOOP across countries provides support for the assumption of LCP—which allows for deviations from the LOOP—in theoretical models. Theoretical models that rely on PCP—which assumes away deviations from the LOOP—fail to take account of important features of international consumer price data. While our results, along with those from the extant literature, provide support for the assumption of LCP with respect to consumer goods, Obstfeld and Rogoff (2000) and Obstfeld (2002) argue that the assumption of PCP is more appropriate at earlier points in the production process. For example, a depreciation of the domestic currency may fail to significantly lower the price of domestically produced consumer goods relative to consumer goods produced abroad because of LCP behavior with respect to consumer goods. Nevertheless, multinational firms may find that domestic labor and other inputs become relatively cheaper, so that they shift some of their production from foreign to domestic plants. This can reopen the expenditure-switching channel closed off by LCP. We view measuring the persistence and volatility in deviations from the LOOP at earlier points in the production process as an important area for future research.

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Table 1: Cross-section regression results for $\hat{\alpha}_{MU}^{i,j}$ and $(\hat{\sigma}_e^{i,j})^2$, Engel and Rogers (1996) Canadian and U.S. city data

(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$\hat{\alpha}_{MU}^{i,j}$ regression			$(\hat{\sigma}_e^{i,j})^2$ regression		
CPI category	Distance coefficient	Border coefficient	R^2	Distance Coefficient	Border coefficient	R^2
All items	0.003	0.031*	0.361	0.027**	1.113**	0.993
	(0.767)	(2.437)		(2.844)	(79.41)	
Food at home	0.000	0.027**	0.362	0.274**	1.612**	0.862
	(0.028)	(4.277)		(3.018)	(22.22)	
Food away from home	0.0088**	0.015**	0.391	0.015	1.266**	0.980
	(2.795)	(4.005)		(1.094)	(53.95)	
Alcoholic beverages	-0.001	0.009†	0.368	0.122†	1.356**	0.894
	(-0.580)	(1.678)		(1.758)	(25.01)	
Shelter	0.007**	0.008**	0.585	0.034**	1.275**	0.993
	(3.858)	(2.658)		(2.715)	(40.07)	
Fuel and other utilities	-0.004	0.036**	0.547	0.515*	3.683**	0.954
	(-1.290)	(3.066)		(2.463)	(8.55)	
Household furnishing and operations	0.007	0.035*	0.322	0.029†	1.291**	0.989
	(1.081)	(2.245)		(1.817)	(42.33)	
Men's and boy's apparel	0.009	0.018*	0.401	0.270*	3.908**	0.990
	(1.626)	(2.270)		(2.461)	(21.80)	
Women's and girl's apparel	0.009	0.013	0.369	0.452**	5.542**	0.996
	(1.278)	(1.024)		(3.848)	(15.10)	
Footwear	0.024	0.0251	0.197	0.063	2.258**	0.990
	(1.526)	(1.279)		(0.366)	(6.389)	
Private transportation	0.013†	0.028**	0.531	0.214**	2.156**	0.941
	(1.654)	(3.211)		(3.654)	(32.74)	
Public transportation	0.002	0.032**	0.548	1.079*	12.19**	0.901
	(0.512)	(7.181)		(2.562)	(16.90)	
Medical care	0.002	0.026**	0.361	-0.016	1.474**	0.989
	(0.320)	(6.348)		(-0.289)	(28.22)	
Personal care	0.006	0.002	0.315	0.029	1.447**	0.979
	(1.400)	(0.312)		(1.138)	(28.71)	
Entertainment	0.012**	0.010†	0.343	0.0587**	1.401**	0.989
	(3.350)	(1.807)		(3.200)	(39.25)	

Notes: heteroskedasticy-consistent *t*-statistics are given in parentheses; †,*,** indicate significance at the 10%, 5%, and 1% significance levels; for columns (5) and (6), the reported values of the coefficients are the actual values multiplied by 10,000.

Table 2: Cross-section regression results for $\hat{\alpha}_{LB}^{i,j}$, Engel and Rogers (1996) Canadian and U.S. city data

(1) CPI category	(2) Distance coefficient	(3) Border coefficient	(4) R^2
All items	0.009 (1.409)	0.056** (3.319)	0.483
Food at home	-0.002 (-0.335)	0.064** (8.937)	0.508
Food away from home	0.014** (3.319)	0.024** (3.962)	0.519
Alcoholic beverages	-0.002 (-0.646)	0.048** (9.472)	0.692
Shelter	0.014** (5.881)	0.012** (4.425)	0.733
Fuel and other utilities	-0.004 (-1.079)	0.064** (3.285)	0.507
Household furnishing and operations	0.007 (0.820)	0.080** (4.048)	0.511
Men’s and boy’s apparel	0.007 (0.720)	0.080** (8.417)	0.584
Women’s and girl’s apparel	0.010 (0.800)	0.070** (3.223)	0.418
Footwear	0.043† (1.670)	0.071* (2.512)	0.248
Private transportation	0.015 (1.382)	0.059** (4.664)	0.634
Public transportation	-0.003 (-0.747)	0.039** (5.926)	0.506
Medical care	0.006 (0.807)	0.049** (8.818)	0.465
Personal care	0.016** (3.029)	0.016 (1.598)	0.421
Entertainment	-0.001 (-0.232)	0.038** (4.641)	0.448

Notes: heteroskedasticy-consistent *t*-statistics are given in parentheses; †,*,** indicate significance at the 10%, 5%, and 1% significance levels; for columns (5) and (6), the reported values of the coefficients are the actual values multiplied by 10,000.

Table 3: Cross-section regression results for $\hat{\alpha}_{MU}^{i,j}$, $\hat{\alpha}_{LB}^{i,j}$, and $(\hat{\sigma}_{e}^{i,j})^2$, Engel and Rogers (2001) European city data

(1) Distance Coefficient	(2) Border coefficient	(3) R^2	(4) Distance Coefficient	(5) Border coefficient	(6) R^2
$\hat{\alpha}_{MU}^{i,j}$ regression			$(\hat{\sigma}_{e}^{i,j})^2$ regression		
0.580	0.009**	0.098	0.127	11.99**	0.675
(0.088)	(4.373)		(1.449)	(56.58)	
$\hat{\alpha}_{LB}^{i,j}$ regression					
4.784	0.022**	0.152			
(0.493)	(7.114)				

Notes: heteroskedasticy-consistent *t*-statistics are given in parentheses; †,*,** indicate significance at the 10%, 5%, and 1% significance levels; the reported coefficient estimates in columns (1), (4), and (5) are the actual values multiplied by 10,000.

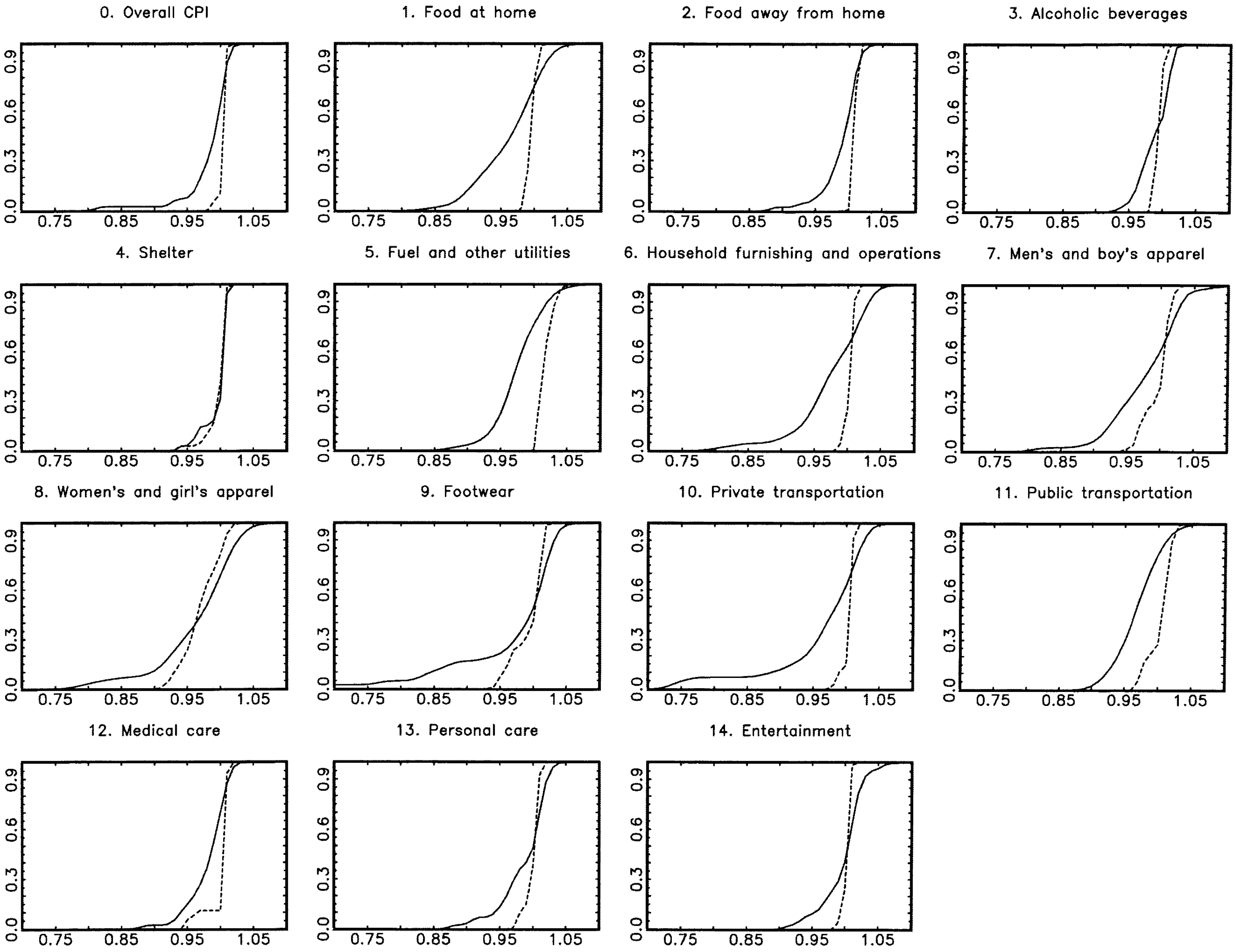
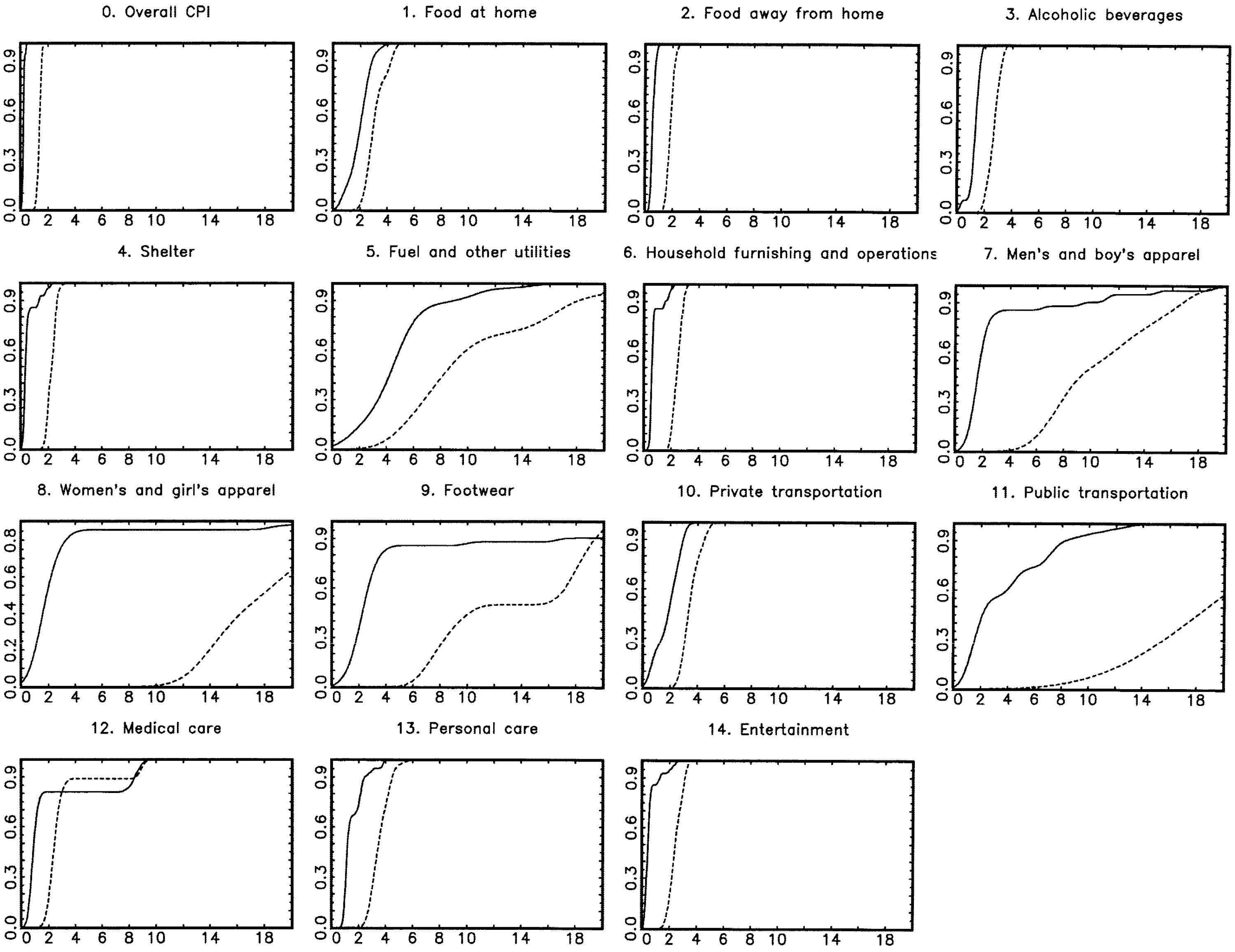

0. Overall CPI
1. Food at home
2. Food away from home
3. Alcoholic beverages
4. Shelter
5. Fuel and other utilities
6. Household furnishing and operations
7. Men's and boy's apparel
8. Women's and girl's apparel
9. Footwear
10. Private transportation
11. Public transportation
12. Medical care
13. Personal care
14. Entertainment

Figure 1: Empirical cumulative density functions for $\hat{\alpha}_{MU}^{i,j}$, Engel and Rogers (1996) Canadian and U.S. city data

Note: solid lines indicate the empirical cumulative density function for city pairs within the same country; dashed lines indicate the empirical cumulative density function for city pairs lying across the border.



0. Overall CPI
1. Food at home
2. Food away from home
3. Alcoholic beverages
4. Shelter
5. Fuel and other utilities
6. Household furnishing and operations
7. Men's and boy's apparel
8. Women's and girl's apparel
9. Footwear
10. Private transportation
11. Public transportation
12. Medical care
13. Personal care
14. Entertainment

Figure 2: Empirical cumulative density functions for $(\hat{\sigma}_e^{i,j})^2$, Engel and Rogers (1996) Canadian and U.S. city data

Note: solid lines indicate the empirical cumulative density function for city pairs within the same country; dashed lines indicate the empirical cumulative density function for city pairs lying across the border.