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E C O N O M I C S   B U L L E T I N

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1 Introduction

Since the 1980s there has been a large number of empirical studies considering the extent to which changes in exchange rates are passed through into import prices. While these studies have used a range of empirical methodologies, most have focused on the industrialized countries. Menon (1995) surveyed 48 studies on exchange rate pass-through and observed that most of the research in this area is done using U.S. or Japanese data. Goldberg and Knetter (1997) note that in the 1980s research on exchange rate pass-through was dominated by the analysis of pass-through to the U.S. More recently, however, some work on exchange rate pass-through has been done for developing countries (see Alba and Papell (1998) and Anaya (2000))[1].

The aim of this paper is to contribute to the analysis of exchange rate pass-through into import prices in developing countries. Developing countries have common characteristics in that they tend to be price takers on international markets and are dependent on imports from industrialized countries. But this does not necessarily imply that the degree of exchange rate pass-through will be the same for all of them. Here we investigate the nature of the long run exchange rate pass-through into import prices in developing countries, and in particular, we try to determine if this long run exchange rate pass-through phenomenon is homogeneous or heterogeneous. In order to do this we define and estimate an import price equation across a panel of 24 developing countries. This equation stipulates that the degree of exchange rate pass-through into import prices is determined by a combination of the nominal effective exchange rate, the prices of competing domestic products, the exporter's costs and domestic demand conditions. We adopt a multi-country framework and use non-stationary panel estimation techniques and tests for panel cointegration. The advantage of using a panel is that the additional information available in the cross-section is a means of increasing the power of tests to identify the absence of spurious cointegration between the variables in our equations, relative to single country tests. No other study has applied a non-stationary panel cointegration and estimation approach in this context.

Our analysis reveals that long run exchange rate pass-through in developing countries is an heterogeneous phenomenon. Our results should provide a deeper understanding of exchange rate pass-through into import prices in developing countries that can be used both for international monetary policy and international trade policy.

This paper is organized as follows. Firstly, we define our price equation. Secondly, we perform the stationarity and cointegration tests. Then, by using the appropriate estimation techniques of our long run relation, we show that the long run exchange rate pass-through in developing countries is heterogeneous. Finally, we provide some concluding remarks.

2 Exchange rate Pass-through equation

The empirical studies of exchange rate pass-through have largely been interested in the extent to which exchange rate movements are transmitted to the pricing of traded goods versus absorbed in producer profit margins or markups. According to Goldberg and Knetter (1997) exchange rate pass-through is defined as the percentage change in the local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries. The studies of exchange rate pass-through into import prices are empirically implemented as a statistical relationship of the elasticity of import prices to exchange rates. Testing this relationship is based on the following

[1] These studies analysed exchange rate passt-hrough to inflation.

equation:

$$\Delta p_t = \gamma \Delta e_t + \varepsilon_t. \tag{1}$$

where p_t and e_t are the natural logarithm of import price and the nominal exchange rate and ε is an error term and γ is the exchange rate pass-through coefficient. The extent of exchange rate pass-through coefficient is based on the value of γ. A one to one response of import prices to exchange rate is known as a complete exchange rate pass-through and $\gamma = 1$, while less than exchange rate pass-through coefficient ($\gamma < 1$) is known as partial or incomplete exchange rate pass-through.

However, Campa and Goldberg (2003) criticize this specification because it only represents a non-structural statistical relationship and lacks an economic interpretation. They argue that a correct specification should include, additionally, controls to capture exporter's costs associated with local inputs and demand conditions in the destination country. Recent empirical studies[2] on exchange rate pass-through into import prices use an approach based on micro-foundations of pricing behavior by exporters firms.

In this paper, the equation that we use to estimate the degree of the exchange rate pass-through into import prices is similar to the equation used in the literature in this area (Hooper and Mann (1989), Goldberg and Knetter (1997) and Campa and Goldberg (2003)). We consider a representative foreign firm having some degree of control over the price of its goods in an importing country. Assume that this representative firm establishes the price of its exports to country i (i is a developing country) in its own currency (PX_{it}) at a markup (λ_{it}) over its marginal cost of production (C^*_{it}), that is:

$$PX_{it} = \lambda_{it} C^*_{it}. \tag{2}$$

The import price in the domestic currency PM_{it} is obtained by multiplying export price PX_{it} by the exchange rate of the importing country i, E_{it}, that is,

$$PM_{it} = E_{it} PX_{it} = E_{it} \lambda_{it} C^*_{it}. \tag{3}$$

The markup is assumed to respond to both demand pressure for exporting country (Y^*_{it}) and competitive pressure in importing country. Competitive pressure in importing country is measured by the gap between the competitor prices in the importing country market (P_{it}) and production cost of exporting firm. Therefore, according to Hooper and Man (1989) the markup λ_{it} is given by

$$\lambda_{it} = \left[\frac{P_{it}}{E_{it} C^*_{it}}\right]^{\alpha} Y^{*\beta}_{it} \quad , 0 < \alpha < 1, \text{ and } 0 < \beta < 1. \tag{4}$$

Substituting equation (4) into equation (2), we obtain

$$PM_{it} = (E_{it} C^*_{it})^{1-\alpha} (P_{it})^{\alpha} Y^{*\beta}_{it}. \tag{5}$$

The logarithmic form of the equation (5) is thus

$$pm_{it} = (1-\alpha) e_{it} + (1-\alpha) c^*_{it} + \alpha p_{it} + \beta y^*_{it}. \tag{6}$$

where lowercase letters denote the logarithmic values of the variables.

In equation (6), the exchange rate pass-through, defined as the partial elasticity of import price with respect to exchange rate, is $(1-\alpha)$. One weakness of this equation is that the pass-through of exchange rate and foreign cost into import price are the same. However, in practice, this restriction

[2] Campa and Goldberg (2003) and Eiji Fuji (2004).

does not necessarily hold. Indeed, Bache (2002) argue that exchange rates are more variable than costs, and a reasonable conjecture is that exporters will be more willing to absorb into their markups changes in exchange rates than change in costs, which are likely to be permanent. Moreover, Athukorala and Menon (1995) have provided purely economic reasons to justify that the coefficient restrictions may not hold such as the incompatibility of price proxies which may result from differences in aggregation level and methods of data collection. Therefore, in estimation, we relax these restrictions and consider the following equation (the long run relationship):

$$pm_{it} = \alpha_i + \beta_1 e_{it} + \beta_2 c^*_{it} + \beta_3 p_{it} + \beta_4 y^*_{it} + \varepsilon_{it}.^3 \tag{7}$$

In this equation, the marginal cost of production of foreign firm is difficult to measure, therefore we adopt the Wholesale price movements of major trade partners of country i (see Eiji Fujii (2004)) represented by

$$C^*_{it} = Q_{it} \times \frac{\widetilde{P}_{it}}{E_{it}}. \tag{8}$$

where E_{it} is the nominal effective exchange rate of country i, $\widetilde{P}_{it}$ is the wholesale price index of country i and Q_{it} is the real effective exchange rate of country i. Taking the logarithm of each variable form, we consider:

$$c^*_{it} = q_{it} - e_{it} + \widetilde{p}_{it}. \tag{9}$$

About the other variables in equation (7), the proxy for domestic competitor's price P_{it} is the Producer Price Index of country i (PPI). As the proxy for the demand pressure Y^*_{it}, we use the GDP of country i and, for import price PM_{it}, we take the import unit value in domestic currency.

3 Data Sources and Empirical Methodology

3.1 Data Sources

The main problem in empirical studies on developing countries is data availability. Because of the difficulty of finding some variables (such as the nominal effective exchange rate), we are only able to consider a panel of 24 developing countries. The data are annual and span the period 1980- 2003 (24 years). They are obtained from International Financial Statistics.

3.2 Panel unit root tests

As a pre-test for cointegration analysis, we first investigate the panel non-stationarity of the variables. Here two types of panel unit root tests are employed: the t-bar test proposed by Im, Pesaran and Shin (2003) (henceforth IPS) and the test proposed by Hadri (2000). The former, a panel analogue of Said and Dickey (1984), tests the null hypothesis of non stationarity, while the latter, a panel analogue of Kwiatkowski et al (KPSS, 1992), tests the null hypothesis of stationarity. The Hadri test has two main advantages when compared with the classical IPS methodology. Firstly, it avoids the lack of power of the unit root-based tests by assuming stationarity under the null hypothesis. Secondly, it is particularly suited for panel data series with short time dimension, which is the case here. When applying the above two tests, an important problem is the cross section dependence (the error terms between the individual errors can be correlated). To deal with this issue, different approaches have been proposed in the literature. Some authors add time dummies to the regressions. Others, like

[3] β_1 is the long-run exchange rate pass-through.

Phillips and Sul (2003), use panel unbiased estimators. One can also remove the "aggregate" effects by subtracting cross-section means from the original observations. In our case, we adopt the last alternative and work with demeaned data.[4]

The IPS test results are shown in Table 1. We compare the observed values to the critical values given in Table 4 of Im, Pesaran and Shin (2003) at the 5% level for N=24 and T=24. We thus conclude that all variables are stationary in first difference. The Hadri test results are shown in Table 2.

Table 1: **IPS panel unit root tests results**

variables	level		first difference	
	intercept	intercept+ trend	intercept	intercept + trend
pim	-1.252	-0.932	-7.420	-7.896
en	-1.489	-2.189	-3.499	-3.105
ppi	-1.523	-0.895	-5.849	-5.915
y*	-1.242	-1.544	-5.016	-5.931
c*	-1.439	-0.955	-3.792	-3.162

Note: : the critical value at the 5% level is -1.73 for the model with an intercept and -2.45 for the model with an intercept and linear time trend.

Table 2: **Hadri panel unit root tests results**

variables	level	first difference
en	4.928	1.329
pim	3.672	0.978
ppi	6.143	-2.234
y*	2.478	-3.765
c*	2.326	0.765

Note: The null of stationarity is rejected if the computed Hadri statistic is greater than 1.645 at the 5% level.

3.3 Tests for panel cointegration

Several authors have recently proposed alternative procedures for panel cointegration tests. In order to ensure robustness of results, we employ Pedroni's tests. Pedroni (1995, 1999) has developed seven tests based on the residuals from the cointegrating panel regression under the null hypothesis of non-stationarity. The first four Pedroni tests are based on the within panel estimator, that are known as the Panel Statistics: and are a variance ration test (v-statistic), a panel version of the Phillips and Perron (1988) ρ-statistic and t-statistic (non-parametric), and the ADF t-statistic (parametric). The additional three statistics are based on pooling along the between dimension and they are known as Group Mean Panel Tests. The three Group Mean statistics are extensions of the Phillips and Perron (1998), ρ-statistic and t-statistic and a parametric t-statistic. As shown in Table 3, all test statistics reject the null of no cointegration.[5]

[4] $\widetilde{y_{it}} = y_{it} - y_{.t}$ where $y_{.t} = \frac{1}{N}\sum_{t=1}^{N} y_{it}$.

[5] Except the v-stat, all test statistics have a critical value of -1.64 (if the test statistic is less than -1.64, we reject the null of no cointegration). The v-stat has a critical value of 1.64 (if the test statistic is greater than 1.64, we reject the null of no cointegration).

* we reject the null of no cointegration.

Table 3: **Pedroni's cointegration tests results**

Statistics	*values*
Panel v-stat	2.438**
Panel rho-stat	-2.881**
Panel pp-stat	-2.532**
Panel adf-stat	-3.238**
group rho-stat	-3.409**
group pp-stat	-3.166**
group adf-stat	-2.304**

4 Long run exchange rate pass-through estimations

4.1 PMG and MG Estimations

Previous empirical work which estimated pass-through elasticities, specified equation (7) in first-differences (Campa and Goldberg (2004) and Bailliu and Fujii (2004)). This type of specification allows estimation of short-run and long-run pass-through. However, in our empirical approach, we need to use a technique that is suitable for dynamic panel data and which allows us to take into consideration non-stationarity of variables and cointegration relationship. To better illustrate this point, we use the « Pooled Mean Group estimator » (PMG) proposed by Pesaran, Shin and Smith (2000). The PMG method restricts the long-run coefficients to be equal over the cross-sections but allows for the short-run coefficients and error variance to differ across groups on the cross-sections. We test for long-run homogeneity using a joint Hausman test[6] based on the null hypothesis of equivalence between the PMG and Mean Group estimator proposed by Pesaran and Smith (1995). The Mean Group estimator is an average of N individual estimations allowing long-run heterogeneity. If we reject the null, we reject the homogeneity of our cross-section's long-run coefficients.

We estimate the following model:

$$\Delta p_{it}^{im} = \phi_i p_{it-1}^{im} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta p_{it-j}^{im} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{it-j} + \mu_i + \varepsilon_{it}. \tag{10}$$

where x_{it} is the vector of explanatory variables : $e_{it}, c_{it}^{*}, p_{it}$ and y_{it}^{*} for country i and μ_i are the fixed effect.

The pooled mean group restriction is that the elements of are common across countries:

$$\Delta p_{it}^{im} = \phi_i p_{it-1}^{im} + \beta^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta p_{it-j}^{im} + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{it-j} + \mu_i + \varepsilon_{it}. \tag{11}$$

In our empirical exploration, we use two different estimations. First, we restrict all long-run coefficients to be equal over the cross-sections and in the second, the homogeneity is imposed only for the long-run pass-through coefficient. In both cases, the Hausman test rejects the assumption of long-run homogeneity.

The PMG and Mean Group estimations for the first case[7] are shown in Table 4.1. The PMG and Mean Group estimation provides significant short run (0.506)[8] and long-run pass-through coefficients

[6] More details will be provided in Appendix.

[7] All PMG and MG estimations were performed using the GAUSS code written by Yongcheol Shin. The program is available on line at http://www.eco.cam.ac.uk/faculty/pesaran/jasa.exe.

[8] Given our data frequency, the short run here refers to one year period.

(respectively 0.637 and 0.726). Secondly, by the joint Hausman test, we reject long-run homogeneity with a probability value of 0.03. For the second case (see table 4.2), we obtain by PMG estimations a short run coefficient of 0.510 and a long run exchange rate pass-through of 0.789. Mean group estimations provide a long run exchange rate pass-through of 0.716. By the Hausman test, we reject the long run homogeneity of exchange rate pass-through coefficient (with a probability value of 0.0056). So, following these results, we conclude that the long run exchange rate pass-through into import prices in developing countries are an heterogeneous phenomenon. Therefore, we now use estimation techniques taking into account the heterogeneity of long-run coefficients.

Table 4.1: **PMG and MG estimations (Homogeneity of all long-run coefficients)**

Estimators	**PMG**		**MG**	
variables	**coefficients**	**t-values**	**coefficients**	**t-values**
en	**0.637**	**7.968**	**0.726**	**2.434**
ppi	**0.449**	**11.859**	**0.237**	**2.722**
y^*	**0.454**	**10.841**	**0.393**	**6.021**
c*	**0.327**	**2.761**	**0.296**	**1.299**

Table 4.2: **PMG and MG estimations (Homogeneity of long-run exchange rate pass-through coefficient)**

Estimators	**PMG**		**MG**	
variables	**coefficients**	**t-values**	**coefficients**	**t-values**
en	**0.799**	**32.471**	**0.716**	**2.731**
ppi	**0.869**	**31.414**	**0.210**	**3.848**
y^*	**0.768**	**5.868**	**0.530**	**8.584**
c*	**0.283**	**6.356**	**0.309**	**9.661**

4.2 Mean Group Panel Estimations

In order to estimate the long run coefficients of the cointegration relationship (7), we use FMOLS and DOLS between-dimension estimators (Group Mean Estimator) as proposed by Pedroni (2001). An important advantage of the between-dimension estimators is that the form in which the data are pooled allows for greater flexibility in the presence of heterogeneous cointegrating vectors. Another advantage is that the estimates have a more useful interpretation when the true cointegrating vectors are heterogeneous. Specifically, the estimates for the between dimension estimator can be interpreted as the mean value for the cointegrating vectors, which is not the case for the within-dimension estimations. By analyzing the results of FMOLS and DOLS estimations, we show that developing countries experience a higher long-run exchange rate pass-through coefficient. With FMOLS, we obtain an estimation of long- run exchange rate pass-through of 77.2% and with DOLS of 82.7% (see Tables 5). However, the pass-through coefficient is not close to one [9]. However, the average masks cross-country difference in long run exchange rate pass-through into import prices. For example, by FMOLS, the long-run pass-through coefficients vary from 107% for Algeria (a complete pass-through coefficient: $\beta_1 > 1$) to 42% for Chile (a partial pass-through coefficient $0 < \beta_1 < 1$) (See Table 6). Similarly, by DOLS the long run pass-through coefficients vary from 110% for Paraguay to 43% for Singapore (See table 7).

Table 5: **FMOLS and DOLS Mean Group Panel estimation**

[9] Campa and Goldberg (2003) find that full pass-through is generally supported as a longer run characterization.

Estimator	FMOLS		DOLS	
variables	coefficients	t-values	coefficients	t-values
en	0.772	2.354	0.827	6.322
ppi	0.243	5.947	0.303	4.178
y*	0.486	2.546	0.920	2.256
c*	0.286	4.178	0.291	2.234

5 Concluding remarks

In this paper we have estimated long-run exchange rate pass-through into import prices equations for a panel of 24 developing countries using a non stationary panel approach. The advantage of using a panel is to try and use the additional information available in the cross-section as a means of increasing the power tests to identify non-spurious cointegration between the variables in our equations relative to single country tests. We have shown that exchange rate pass-through is determined by a combination of the nominal effective exchange rate, the price of competing domestic product, the exporter's cost and domestic demand conditions. We used the panel cointegration test of Pedroni (1995, 1999) to show that there is some evidence of cointegration between all variables. We estimated our long-run equation using several panel estimators, namely PMG, MG, FMOLS and DOLS. By estimating our long-run relationship using PMGE and MGE approach, we find heterogeneity in the long-run exchange rate pass-through into import prices in developing countries. Then, by using FMOLS and DOLS between-dimension estimators (Group Mean Estimator) proposed by Pedroni (2001) we find considerable heterogeneity of long-run coefficients, in particular long-run exchange rate pass-through coefficients. We find that these countries experience on average a high long run exchange rate pass-through (by FMOLS, we obtain 77.25% and by DOLS, we obtain 82.7%). The direction for future research follows naturally - to analyze the determinants of the differences in exchange rate pass-through in developing countries.

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Table 6: **FMOLS estimations by country**

country	en	ppi	y*	c*
1-Algeria	1.07 (20.15)	-0.27 (-2.37)	0.33 (2.27)	-0.40 (-2.77)
2-Burkina Fasso	0.56 (3.25)	-0.35 (-0.42)	-0.68(-1.34)	-1.13(-2.26)
3-Botswana	0.37 (4.83)	0.57 (5.54)	-0.42(-1.86)	-1.26(-3.31)
4-Cote Ivoire	0.73 (15.63)	-0.28 (-1.15)	0.02(0.05)	0.03(4.18)
5-Gabon	0.43 (2.45)	0.86 (9.03)	-0.82(-4.11)	0.22(2.71)
6-Moroco	0.93 (6.20)	0.73 (1.11)	-0.17(-0.31)	1.65(1.68)
7-Nigeria	0.64 (5.17)	0.77 (2.67)	0.78 (1.10)	-1.20 (-3.06)
8-Senegal	1.11 (2.71)	-0.10 (-1.69)	1.14(3.67)	1.71 (5.44)
9-Tunisia	0.33 (3.02)	0.23 (1.10)	0.02 (10.02)	-0.31 (-0.74)
10-Zambia	0.88 (10.93)	-0.15 (-11.79)	1.69 (3.24)	1.55 (4.85)
11-India	0.55 (3.03)	1.34 (5.89)	3.60 (0.90)	-1.19 (-4.98)
12-Indonesia	0.29 (2.10)	0.62 (1.39)	-0.84 (-2.57)	-1.08 (-1.67)
13-Iran	0.27 (1.55)	0.51 (5.68)	-1.15 (-2.08)	0.98 (3.39)
14-Pakistan	0.47 (4.29)	0.48 (0.98)	0.07 (0.45)	-0.13 (-3.32)
15-Phillipines	0.68 (11.41)	0.80 (2.01)	0.18 (0.21)	0.31 (0.93)
16-Singapour	0.65 (3.69)	0.95 (2.42)	2.32 (2.05)	2.22 (0.77)
17-Bolivia	1.17 (3.64)	1.06 (4.82)	-0.08 (-0.19)	2.29 (2.41)
18-Chile	0.42 (6.07)	-0.34 (-5.57)	0.43 (2.45)	0.10(0.76)
19-Colombia	0.74 (4.85)	4.19 (10.14)	1.52 (2.46)	2.73 (2.83)
20-Costa Rica	2.03 (0.91)	-0.29(-0.19)	0.68 (0.18)	0.87 (1.24)
21-Equador	1.21 (1.16)	-1.59(-1.32)	7.87 (3.22)	1.38 (3.02)
22-Paraguay	0.95 (2.69)	-2.10 (-4.98)	-0.18 (-1.06)	2.72 (6.91)
23-Uruguay	1.02 (3.98)	0.38 (2.76)	-1.84 (-6.36)	0.05(0.32)
24-Venezuela	1.03 (2.82)	-2.17 (-5.10)	-0.17 (-0.99)	1.14 (3.27)

Table 7: **DOLS estimations by country**

country	en	ppi	y*	c*
1-Algeria	1.34 (0.87)	0.16 (0.29)	2.52 (8.29)	0.11 (1.20)
2-Burkina Fasso	0.46 (2.66)	0.39 (0.53)	0.32 (1.76)	4.50 (9.01)
3-Botswana	0.50 (2.22)	0.72 (2.11)	0.42 (2.14)	-0.44 (1.12)
4-Cote Ivoire	1.03 (3.99)	1.44 (2.89)	0.05 (0.16)	-2.71 (-2.72)
5-Gabon	0.39 (2.76)	0.85 (2.43)	0.12 (0.89)	0.565 (0.09)
6-Moroco	1.12 (3.67)	0.95 (3.19)	0.13 (0.31)	-0.15 (-4.71)
7-Nigeria	0.45 (1.68)	0.61 (0.39)	0.23 (0.42)	2.42 (0.20)
8-Senegal	1.12 (2.54)	0.76 (1.87)	-0.83 (-3.24)	0.23 (1.79)
9-Tunisia	0.39 (3.55)	0.63 (0.26)	0.12 (0.58)	0.33 (1.11)
10-Zambia	0.42 (2.48)	-0.72 (-0.86)	1.99 (1.39)	-1.40 (-6.71)
11-India	0.97 (2.42)	-0.51 (-2.08)	1.32 (2.01)	-5.64(-0.43)
12-Indonesia	0.41 (10.38)	0.58 (0.33)	-0.14 (-0.52)	-0.50 (-3.12)
13-Iran	0.37 (1.14)	-6.14 (-5.66)	5.86 (5.42)	3.59 (1.96)
14-Pakistan	0.43 (2.37)	0.94 (3.33)	-0.14 (-2.52)	-0.12 (-0.84)
15-Phillipines	0.75 (2.11)	2.88 (1.24)	3.75 (1.61)	0.97 (2.13)
16-Singapour	0.43 (2.08)	0.68 (2.31)	0.34 (1.82)	-1.16 (-1.94)
17-Bolivia	1.63 (1.26)	0.99 (0.13)	-0.14 (-0.12)	1.21 (1.06)
18-Chili	0.42 (2.99)	-0.24 (-0.96)	1.36 (1.16)	0.14 (0.18)
19-Colombia	0.67 (4.70)	1.43 (0.97)	-0.01 (-0.70)	0.26 (1.03)
20-Costa Rica	2.09 (0.44)	-0.60 (-0.30)	1.73 (0.33)	0.42 (0.59)
21-Equator	1.03 (2.30)	1.25 (1.90)	0.31 (1.09)	0.80 (1.87)
22-Paragay	1.10 (3.06)	0.08 (11.06)	0.84 (16.05)	1.67 (0.85)
23-Urugay	0.95 (4.25)	0.07 (2.06)	0.36 (1.41)	0.17 (0.19)
24-Venezuela	1.29 (1.09)	-0.01 (-0.24)	1.54 (0.07)	1.09 (0.93)