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The Big Mac Standard: A statistical Illustration

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Abstract

We demonstrate a statistical procedure for selecting the most suitable empirical model to test an economic theory, using the example of the test for purchasing power parity based on the Big Mac Index. Our results show that supporting evidence for purchasing power parity, conditional on the Balassa−Samuelson effect, depends crucially on the selection of models, sample periods and economies used for estimations.

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1. Introduction

The well-known law of one price, one of the strongest versions of purchasing power parity (PPP), requires, for any tradable goods *i*, $p_i = Ep_i^*$, where p_i is the nominal price of good i in the domestic currency, E is the domestic price of foreign currency, and p_i^* is the nominal price of good i in foreign currency.

Following Nelson (2001), we briefly summarize the empirical literature testing purchasing power parity using time series macroeconomic data.

Before the 1980s, researchers testing purchasing power parity usually estimated equation (1) and tested the joint hypothesis of $\alpha = 0,\ \beta = 1$,

$$\ln(E_t) = \alpha + \beta \cdot [\ln(P_t) - \ln(P_t^*)] + \varepsilon_t, \tag{1}$$

where E_t is the nominal exchange rate, P_t is the domestic price index, P_t^* is the foreign price index, ε_t is an error term, and subscript t indicates time t.

Since Dickey and Fuller (1979, 1981) proposed formal tests for the existence of a unit root, it has been reported that most macroeconomic time series, including nominal exchange rate and price level, had a unit root. Given recent developments econometric analysis of non-stationary data, a test for purchasing power parity typically tests whether the real exchange rate, or $q_t = \ln(E_t) - \ln(P_t) + \ln(P_t^*)$, is stationary or not. A major problem in such lines of research is that the test for purchasing power parity requires a very long run of data to reject the null hypothesis of the existence of a unit root. Thus, recently, economists have emphasized the use of unit root tests that apply to panel data (see for an earlier example, Wu, 1993).

According to Nelson (2001), such studies find that purchasing power parity tends to hold if researchers use long-run time series data. In addition, the half-life of the deviation from purchasing power parity is, on average, 3.7 years when one uses quarterly CPI data and the nominal exchange rate relative to the US dollar (Nelson, 2001, Table 7.2). One of the most promising reasons for such deviations from purchasing power parity is the Balassa–Samuelson effect (Balassa, 1963; Samuelson, 1964). The other hypothesis is that the firm's response to changes in nominal prices does not follow immediately if shocks in the nominal exchange rate are perceived to be temporary.

Most empirical studies on purchasing power parity have employed aggregated data, since it is not easy to design a common basket of commodities to make consistent cross-country comparison of nominal price levels. In this context, the annual publication of the Big Mac Index by *The Economist* is a notable exception. Essentially, the Big Mac PPP is the exchange rate that would mean hamburgers cost the same in the US as abroad. Namely, the Big Mac Index tests whether the relative prices of an identical basket of goods and services measured by a McDonald's Big Mac, in terms of domestic currencies, is equal to nominal exchange rates in the financial markets in the long run. This is possible because the Big Mac is produced in about 120 countries.

There are several studies using this Index. For example, Pakko and Pollard (1996) admit that the Big Mac Index is useful for testing purchasing power parity. However, they conclude that purchasing power parity does not hold for the following reasons: barriers to trade, the inclusion of non-traded goods in the Big Mac, imperfect competition in goods markets and factor markets, and current account imbalances. On the other hand, Ong (1997) reports that the Big Max Index is surprisingly accurate in tracking exchange rates over the

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long-term. Click (1996) also finds that purchasing power parity based on the Big Mac Index holds in the time-series dimensions, and that the country-specific deviations from purchasing power parity are explained by the Balassa–Samuelson effect.

In this paper, we show that the results for the acceptance or rejection of purchasing power parity using the Big Max Index are sensitive to the choice of statistical models, and thus it might be desirable to employ various statistical techniques. This result is based not only on statistical tests, but also on the properties of the data under study. More specifically, our contributions in this paper are: (i) to expand data sets up to the year 2002, (ii) to pay attention to the outliers in the data sets, (iii) to introduce estimates using a dynamic panel data model.

The organization of the rest of this paper is as follows. Section 2 discusses our theoretical and empirical framework. Section 3 shows the data sets used in this paper, and section 4 summarizes our empirical findings. Section 5 concludes.

2. Model

Click (1996) estimates equation (2) using the data published in *The Economist* and *The World Development Report*,

$$\ln(\frac{P_{it}}{P_t}) = \alpha + \beta 1 \cdot \ln(E_{it}) + \beta 2 \cdot \ln(\frac{RGDP_{it}}{RGDP_t}) + \varepsilon_{it} \quad (2)$$

where p_{it} is the price of a Big Mac in economy i in local currency, p_t is the price of a Big Mac in the United States, E_{it} is the nominal exchange rate of economy i against the US dollar, $RGDP_{it}$ is real GDP per capita in economy i, $RGDP_t$ is the US real GDP per capita, subscript i indicates country i, and subscript t indicates time t. Purchasing power parity provides us with the null hypothesis that $\alpha = 0,\ \beta 1 = 1,\ \beta 2 = 0$.

Click (1996) reports that, restricting $\beta 2 = 0$, purchasing power parity does not hold for the sample period from 1986 to 1995, as tested by the estimation of equation (2), which uses a pooling model. However, purchasing power parity holds when tested by a one-way fixed-effects model and by a one-way random-effects model, where the one-way fixed-effects model is

$$\ln(\frac{P_{it}}{P_t}) = \alpha(i)_{of} + \beta 1_{of} \cdot \ln(E_{it}) + \beta 2_{of} \cdot \ln(\frac{RGDP_{it}}{RGDP_t}) + \varepsilon_{it}, \quad (3)$$

$$i = 1,\ldots,N, t = 1,\ldots T., E(\varepsilon_{it}) = 0,\ Var(\varepsilon_{it}) = \sigma_e^{\ 2},$$

and the one-way random-effects model is

$$\ln(\frac{P_{it}}{P_t}) = \alpha(0) + \alpha_{or}(i) + \beta 1_{or} \cdot \ln(E_{it}) + \beta 2_{or} \cdot \ln(\frac{RGDP_{it}}{RGDP_t}) + \varepsilon_{it}, \quad (4)$$

$$i = 1,\ldots,N, t = 1,\ldots T., E(\varepsilon_{it}) = 0,\ Var(\varepsilon_{it}) = \sigma_e^{\ 2},$$

$$E(\alpha_{or}(i)) = 0,\ Var(\alpha_{or}(i)) = \sigma_\alpha^{\ 2}.$$

Click (1996) reports that without restricting $\beta 2 = 0$, the pooling model again rejects purchasing power parity, but the one-way random-effects model does not reject purchasing power parity using the sample period 1986 to 1993.

Based on those results, Click (1996) concludes that in the time-series dimension, using the random-effects model, purchasing power parity holds, conditional upon the Balassa–Samuelson effect. His findings are not consistent with those of prior research; it is difficult to support purchasing power parity using short-run time series data. However, other studies using the Big Mac Index often support purchasing power parity (e.g., Ong, 1997). Froot and Rogoff (1996) pointed out that those results could reflect the following reasons: most countries under study officially fixed their nominal exchange rate; some countries in the sample experienced hyperinflation during the periods of estimation; and McDonald's quickly adjusts the price of the Big Mac.

In this paper, we use a statistical test to choose the most suitable empirical model for testing purchasing power parity, and examine whether purchasing power parity holds, following Fujiki and Kitamura (1995).

First, we compare the pooling model with the fixed-effects model, based on the F-test (see Greene, 2000, for details). We also compare the fixed-effects model and the random-effects model, using the Hausman test. Under the null hypothesis $E(\alpha_{or}(i)|X_{it}) = 0$, where X is the explanatory variable in the regression, $(\hat{\beta}_{or} - \hat{\beta}_{of})'[Var(\hat{\beta}_{or} - \hat{\beta}_{of})]^{-1}(\hat{\beta}_{or} - \hat{\beta}_{of})$ follows a χ^2 distribution whose degree of freedom is equal to Dim($\hat{\beta}_{or}$). The decision rule is that if the Hausman test statistic is significantly large, we should reject $\hat{\beta}_{or}$ and accept $\hat{\beta}_{of}$. We can also use the standard Lagrange Multiplier (LM) test to compare the pooling model with the random-effects model. The decision rule is that the larger the LM test statistic, the more likely is rejection of the pooling model and acceptance of the random-effects model.

Second, we incorporate the suggestion made by Froot and Rogoff (1996), and drop an economy from our sample if that economy experiences changes in exchange rate regimes, which usually results in hyperinflation.

Third, Froot and Rogoff (1996) conjectured that McDonald's adjustment of price depends on the nature of the shock, and in turn affects the results of the purchasing power parity test. This point is quite compelling in the literature on pricing to the market. Without large enough shocks, firms do not change their prices quickly. To test this idea, suppose that the desired level of relative price considered by McDonald's, $\ln(P_{it}^{*} / P_t^{*})$, can be approximated as equation (5):

$$\ln(\frac{P_{it}^{*}}{P_t^{*}}) = \alpha(0)^{*} + \beta 1^{*} \cdot \ln(E_{it}) + \beta 2^{*} \cdot \ln(\frac{RGDP_{it}}{RGDP_t}) + \varepsilon_{it}^{*} \ . \tag{5}$$

However, suppose that actual adjustment of price follows a linear mechanism, such that price is adjusted to maintain some constant proportionate gap between desired and actual price, as shown in the following equation:

$$\ln(\frac{P_{it}}{P_t}) - \ln(\frac{P_{it-1}}{P_{t-1}}) = \gamma^{*} \cdot \{\ln(\frac{P_{it}^{*}}{P_t^{*}}) - \ln(\frac{P_{it-1}}{P_{t-1}})\} + u_{it} \ . \tag{6}$$

Then, we could compute the speed of adjustment by running equation (7):

$$\ln(\frac{P_{it}}{P_t}) = \alpha(0)_{DP} + \beta 1_{DP} \cdot \ln(E_{it}) + \beta 2_{DP} \cdot \ln(\frac{RGDP_{it}}{RGDP_t}) + \gamma \cdot \ln(\frac{P_{it-1}}{P_{t-1}}) + \varepsilon_{it} \ , \tag{7}$$

where $\alpha(0)_{DP} = \alpha(0)^{*}\gamma^{*}$, $\beta(1)_{DP} = \beta(1)^{*}\gamma^{*}$, $\beta(2)_{DP} = \beta(2)^{*}\gamma^{*}$, $u_{it} = u_{it} + \varepsilon_{it}^{*}\gamma^{*}$, $\gamma = 1 - \gamma^{*}$. One can define the long-run elasticity of McDonald's pricing with respect to market exchange rate as $\beta 1_{DP}/(1-\gamma)$. Depending on the assumption whether $\alpha(0)_{DP}$ is fixed or random, a fixed-effects (FE) model or a random-effects (RE) model is used to estimate equation (7).

A standard method applicable to a dynamic panel data model is the instrumental variable (IV) or generalized method of moments estimator (GMM) of equation (7).[1] However, Monte Carlo studies conducted by Hsiao, Pesaran, and Tahmiscioglu (2002) show that those standard estimators are subject to serious bias and size distortion in a finite sample, in particular, if γ is close to one. To check robustness, we also provide a minimum distance estimator (MDE) for the fixed-effects model, and generalized least squares (GLS) supplemented with the initial value problem suggested by Hsiao (1986) for the random effects model. Fujiki, Hsiao and Shen (2002) briefly review the merits of using MDE and GLS in their statistical appendix.

3. Data

Following Click (1996), we use the data published in *The Economist* (1986–2002, various issues) for the price of a Big Mac and the nominal exchange rate. We obtain data for relative income from shares of aggregate GDP based on purchasing power parity, the basis for the country weights used to generate the *World Economic Outlook, April 2003*, published by the International Monetary Fund.[2] The data are expressed as a percentage of the world total. Thus, the ratio for the US share divided by the ratio of population to the US will replicate purchasing power parity base relative income per capita. Population figures are also available from the *World Economic Outlook, April 2003*.

Unfortunately, the panel data set provided by *The Economist* is not an ideal balanced panel data set. As Table 1 shows, we survey 34 economies between the year 1986 and 2002, yielding 406 observations. However, only Britain was surveyed in all surveys. Some economies are only rarely surveyed. For example, Greece, Portugal, and Venezuela were surveyed only once, and thus are omitted from the analysis. We also omit the data on Yugoslavia since we have no data on relative GDP. Moreover, since the 2002 survey, the Euro-area economies, France, Germany and Spain, were not surveyed.

If one focuses the analysis between the year 1996 and 2002, we have a balanced panel of 23 economies: Argentine, Australia, Brazil, Britain, Canada, Chile, China, Denmark, Hong Kong, Hungary, Japan, Malaysia, Mexico, New Zealand, Poland, Russia, Singapore, South Africa, South Korea, Sweden, Switzerland, Taiwan, and Thailand. Thus, it is interesting to investigate the results using only that sub-sample of economies. Figure 1 plots all observations from 1996 to 2002 for those 23 economies. It is tempting to conclude that purchasing power parity holds, since most observations are on the 45-degree line. However, as Figure 2 and Figure 3 show, an economy's movement over time might not always lie above the 45-degree line, and thus one might well consider that a fixed-effects estimator that emphasizes time series variation within an economy could give us dramatically different results. Moreover, the data on Russia look very unstable before and after the Russian currency crisis in 1998. Argentina's data shows an unusual jump in 2002, since it abandoned its currency board. One might expect that in such a situation, a one-way fixed-effects model will give us smaller estimates compared with a pooling model.

[1] See for example, Ahn and Schmidt (1995), and Arellano and Bover (1995).

[2] All datasets are downloaded from http://www.imf.org/external/pubs/ft/weo/2003/01/data/index.htm.

4. Results of Regressions

4.1 Results using unbalanced panel data sets

The first three rows of Table 2 report our estimation of equations (2), (3) and (4) using the sample period of 1986–2002 holding $\beta 2 = 0$. The pooling model, model (1), yields β 1=0.99312 and α=-0.0606. The F-value to test the joint null hypothesis of $\alpha = 0,\ \beta 1 = 1$ is 6.5669, and its p-value is 0.0016. Therefore, the pooling model rejects purchasing power parity. However, the F-value for the one-way fixed-effects model, model (2), which tests the one-way fixed-effects model versus the null hypothesis of the pooling model, is 34.117; therefore the one-way fixed-effects model, which reports β1=0.96364, dominates the pooling model. The one-way random-effects model, model (3), shows β1=0.9677 and α=-0.00981. Note that the LM test statistics and the Hausman test statistics suggest that the relevant statistical model is the one-way random-effects model, because the LM test statistic (1215.74, p-value 0.0000) are large enough to reject the null hypothesis of the pooling model against the one-way random-effects model, while the Hausman test statistic (almost 0, p-value 0.9688) is too small to reject the null hypothesis of the one-way random-effects model against the one-way fixed-effects model. The standard error, β1, for the one-way random-effects model is so small that it is possible to reject the null hypothesis that β1=1, while the large standard error of α implies that α=0 cannot be rejected. In sum, we cannot support purchasing power parity once we impose the restriction that β2=0, based on models (1) through (3).

The fourth, fifth and sixth rows of Table 2, models (4) to (6), report the results of estimating equations (2), (3) and (4). The test statistics for the F test, the LM test, and the Hausman test suggest that the relevant statistical model is the one-way random-effects model, model (6). This is because the F-value for the one-way fixed-effects model supports the one-way fixed-effects model against the pooling model, while the LM test statistics are large enough to reject the null hypothesis of the pooling model against the one-way random-effects model. Hausman test statistic for the one-way random-effects model shows one-way random-effects model dominates. Again, model (6) shows a rejection of the null hypothesis that β1=1. Those results are robust to the omission of data on Russia, as can be seen from the seventh to the twelfth rows of Table 2, models (7) to (12), while Figure 1 gives us an impression that the results could be sensitive to Russian outliers.

To replicate the results of Click (1996) that used data up to 1995 for equation (2), and up to 1993 for equation (3) and (4), we estimate models (13) through (18) in Table 1. Contrary to Click (1996), model (13) in Table 1 shows that the joint null hypothesis that $\alpha = 0,\ \beta 1 = 1$, is not rejected at the five percent level, because the value of the F-statistic is only 2.8987 (p-value = 0.0575). However, model (16) shows that once we add the relative per capita income ratio, we reject the null hypothesis of purchasing power parity, that $\alpha = 0, \beta 1 = 1$, because the F-statistic takes the value 9.4223 (p-value = 0.0002). Moreover, inspection of the test statistics for the F-test, the LM test, and the Hausman test suggest that relevant specifications are (15) and (18), as both show β1=0.9621 or 0.98857, whose standard errors are 0.01 and 0.02, respectively. Therefore, over-all, the results shown in Table 2 replicate the results of Click (1996) that purchasing power parity conditional upon the Balassa–Samuelson effect works only in a limited case, (model (18)). Our different results could be explained by different estimates of purchasing power parity base income per capita made since the study of Click (1996).

4.2 Results using balanced panel data sets

4.2.1 Basic Results

Our estimates of equations (2), (3) and (4) using balanced panel data are summarized in Table 3. Inspection of Table 3 shows that, based on the pooling model, models (19) and (22) seem to support purchasing power parity. However, in both models, the null hypothesis of purchasing power parity, that $\alpha = 0, \beta 1 = 1$, is rejected because F-statistics take the value of 39.5854 and 45.6081. Indeed, the statistically preferred model in this table is the random-effects model where the income ratio variable is omitted (model (21)), while including the log of income ratio variable results in the statistically preferred model being the fixed-effects model (24). It is interesting that the coefficients $\beta 1$ are very close to one, irrespective of specifications, and with or without income ratio variables.

4.2.1 Russian Outliers?

Following the suggestion of Froot and Rogoff (1996), that an economy hit by hyperinflation could lead to evidence favorable to purchasing power parity, we drop observations for Russia from the dataset. The results are summarized as models (25) to (30) in Table 3. Observe that models (25) and (28) still show quite similar results to models (19) and (22). However, regarding the fixed-effects model and random-effects model, the value of the coefficient of $\beta 1$ falls substantially from one, and apparently rejects the null hypothesis of purchasing power parity. Therefore, the results using this small balanced panel seem to be sensitive to the inclusion of outliers, especially when one employs either the fixed-effects model or the random-effects model.

4.2.3 Exchange Rate Regimes

The previous section shows that the results of using this small panel might be sensitive to the inclusion of outliers. One could easily imagine that large jumps, or outliers, in the exchange rate data should happen when speculative attacks occur. In this context, it is useful to see whether there are changes in exchange rate regimes during the sample period. Table 4 reports the classification codes of exchange rate regimes proposed by Reinhart and Rogoff (2002). Shaded areas in the table indicate the samples which are included for the analysis of the balanced panel data set. As we can see, we could exclude Brazil, Malaysia, and Korea, as well as Thailand and Russia, since those economies experienced currency crises and thus the codes are renumbered between the years 1996 and 2001. Reinhart's and Rogoff's (2002) data cover up to 2001, and thus we also exclude Argentina. We end up with seven years of data on 17 economies, in total, 119 observations.

Models (31) to (36) shown in Table 3 suggest that a reasonable model for this particular sample is the one-way fixed-effects model with or without the relative GDP ratio, and the estimates of coefficients on market exchange rate are far below one. Overall, evidence shows that while the pooling model seems to be robust to the choice of sample economies, the other two models are not.

4.2.4 McDonald's Pricing Behavior?

It is true that, given the income ratio, models (23), (29) and (35) in Table 3 show that the relative price of a Big Mac with respect to the US price level is relatively expensive in richer economies than the market nominal exchange rate suggests. Hence, one may argue that the Balassa–Samuelson effect might not be the sole reason for departure from purchasing power parity. For example, one could interpret the results as a deliberate price setting behavior for the Big Mac by McDonald's, as Froot and Rogoff (1996) suggested.

To check this idea indirectly, one can estimate a dynamic panel data model and infer the dynamic adjustment mechanism. Table 5 summarizes the results for equation (7). As can be seen from the first to the sixth rows of Table 5 (models (37) to (40)), the estimated parameter values of γ, responses from the lagged dependent variables, are small positive values and take at most 0.3 irrespective of the choice of estimation methods, and the coefficient on the current market exchange rate, $\beta 1$, is still close to one, given the relative GDP ratio. MDE might work better than the IV or GMM in a finite sample; however, since γ is at most 0.3 in model (37), there is no reason to believe that model (38) is superior to model (37). Moreover, long-run elasticity of McDonald's pricing with respect to the market exchange rate, $\beta 1/(1-\gamma)$, is also close to one, except for model (37). These results are consistent with the idea that McDonald's responds to fluctuations of nominal exchange rates quickly, and almost one-to-one.

However, once we drop the data on Russia (models (41) to (44)), or use the samples of stable exchange rate economies (models (45) to (48)), the estimates of $\beta 1$ take values close to zero, and are statistically insignificant except for GLS (models (44) and (48)). Moreover, γ becomes close to one and in some case even exceeds one (models (41) and (42)). Therefore, the application of dynamic panel data models to this particular small data set seems to be risky, especially because it seems to be sensitive to the inclusion of outliers. Thus, it is premature to conclude that dynamic panel data models provide strong evidence for rapid adjustment of sale price by McDonald's.

5. Summary and Conclusion

Evidence that purchasing power parity holds, conditional on the Balassa–Samuelson effect, as reported in Click (1996), is not robust to the choice of methods of estimation, sample economy and sample periods. In addition, one should be careful about the measurement problems inherent in the Big Mac index, which we ignore in this paper. For example, as the index is created from data collected only at one time in a year, that time might coincide with transitory exchange rate fluctuations that do not reflect exchange rates throughout a year.

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Table 1 Data Set

Country	2002	2001	2000	1999	1998	1997	1996	1995	1994	1993	1992	1991	1990	1989	1988	1987	1986	Total
Argentine	1	1	1	1	1	1	1	1	1	1	1							11
Australia	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	16
Austria					1	1	1	1	1									5
Belgium					1	1	1	1	1	1	1	1	1	1	1	1	1	13
Brazil	1	1	1	1	1	1	1	1	1	1	1						1	12
Britain	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	17
Canada	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	16
Chile	1	1	1	1	1	1	1	1	1									9
China	1	1	1	1	1	1	1	1	1	1	1							11
Czech	1	1	1		1	1	1	1	1									8
Denmark	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		16
France		1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	16
Germany		1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	16
HongKong	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	16
Hungary	1	1	1	1	1	1	1	1	1	1	1	1						12
Indonesia	1	1	1	1	1			1										6
Israel	1		1	1	1	1	1	1										7
Ireland										1	1	1	1	1	1	1	1	8
Italy		1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		15
Japan	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	16
Malaysia	1	1	1	1	1	1	1	1	1	1								10
Mexico	1	1	1	1	1	1	1	1	1	1								10
Netherland				1	1	1	1	1	1	1	1	1	1	1	1	1	1	14
NewZealand	1	1	1	1	1	1	1	1										8
Poland	1	1	1	1	1	1	1	1	1									9
Russia	1	1	1	1	1	1	1	1	1	1	1	1	1					13
Singapore	1	1	1	1	1	1	1	1	1		1	1	1	1	1		1	15
SouthAfrica	1	1	1	1	1	1	1											7
SouthKorea	1	1	1	1	1	1	1	1	1	1	1	1	1	1				14
Spain		1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	15
Sweden	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1	16
Switzerland	1	1	1	1	1	1	1	1	1	1								10
Taiwan	1	1	1	1	1	1	1	1	1									9
Thailand	1	1	1	1	1	1	1	1	1	1								10
Total	26	29	30	30	33	32	32	32	29	24	21	18	17	16	15	8	14	406

Note: Omitted economies are: Euro area, Greece, Philippines, Portugal, Peru, Turkey, Venezuela, Yugoslavia.

Table 2 Results of Regressions

Dependent Variable = Ln(P(i,t)/P(t))									
Sample	Model	Methods	α	$\beta 1$	$\beta 2$	$\overline{R}^2$	F (p-value)	LM (p-value)	Hausman (p-value)
1986-2002	(1)	Pool	-0.05060	0.99312		0.9744	6.5669		
(N=405)		(S.E.)	(0.0282)	(0.0080)			(0.0016)		
	(2)	Fixed		0.96364		0.99312	34.117		
		(S.E.)		(0.1047)			(0.0000)		
	(3)	Random	-0.00981	0.9677		0.9735		1215.74	0.00
		(S.E.)	(0.0642)	(0.0103)				(0.0000)	(0.9688)
1986-2002	(4)	Pool	-1.63913	1.01746	0.38821	0.9820	90.8853		
(N=405)		(S.E.)	(0.1246)	(0.0070)	(0.0299)		(0.0000)		
	(5)	Fixed		0.96994	0.39374	0.99324	21.341		
		(S.E.)		(0.1019)	(3.5213)		(0.0000)		
	(6)	Random	-1.4711	0.9818	0.3704	0.9735		1,043.15	0.01
		(S.E.)	(0.2765)	(0.0100)	(0.0690)			(0.0000)	(0.9933)
1986-2002	(7)	Pool	-0.05625	0.99946		0.97443	4.4837		
(N=393)		(S.E.)	(0.0279)	(0.0082)			(0.0141)		
Drop	(8)	Fixed		0.95937		0.99406	41.27400		
Russia		(S.E.)		(0.1455)			(0.0000)		
	(9)	Random	0.00293	0.96538		0.97263		1,267.68	0.00
		(S.E.)	(0.0652)	(0.0110)				(0.0000)	(0.9668)
1986-2002	(10)	Pool	-1.60412	1.02102	0.37862	0.98204	86.7188		
(N=393)		(S.E.)	(0.1233)	(0.0070)	(0.0296)		(0.0000)		
Drop	(11)	Fixed		0.96178	0.32747	0.99412	26.20200		
Russia		(S.E.)		(0.1462)	(4.1161)		(0.0000)		
	(12)	Random	-1.35875	0.97687	0.34497	0.97322		1,161.13	0.01
		(S.E.)	(0.2809)	(0.0105)	(0.0702)			(0.0000)	(0.9945)
1986-1995	(13)	Pool	0.05851	1.00358		0.97664	2.8987		
(N=193)		(S.E.)	(0.0388)	(0.0111)			(0.0575)		
	(14)	Fixed		0.95378		0.99426	19.66000		
		(S.E.)		(0.1597)			(0.0000)		
	(15)	Random	0.11539	0.96217		0.97263		299.43	0.00
		(S.E.)	(0.0704)	(0.0132)				(0.0000)	(0.9589)
1986-1993	(16)	Pool	-1.31523	1.03332	0.32181	0.97931	9.4223		
(N=132)		(S.E.)	(0.3155)	(0.0133)	(0.0739)		(0.0002)		
	(17)	Fixed		0.94885	0.53617	0.99266	10.82400		
		(S.E.)		(0.5083)	(10.3464)		(0.0000)		
	(18)	Random	-1.21152	0.98857	0.32476	0.97322		126.50	0.01
		(S.E.)	(0.4288)	(0.0204)	(0.1040)			(0.0000)	(0.9945)

Notes: F value for the pooling model is the test statistic for the joint null hypothesis that $\alpha = 0,\ \beta 1 = 1$.

F-value for the fixed-effects model compares the fixed-effects model versus the null hypothesis of the pooling model. High values of LM favor the fixed-effects model and the random-effects model over the pooling model. High (low) values of the Hausman test statistic favor the fixed-effects model (random-effects model).

All estimations include data on Russia except in the year 1990, where the income ratio data are missing. Thus, total observations are 405 in the case for equations (1) thought (3), rather than 406 shown in table 1.

Table 3 Results of Regressions

Dependent Variable = Ln(P(i,t)/P(t))									
Sample	Model	Methods	α	$\beta 1$	$\beta 2$	$\overline{R}^2$	F (p-value)	LM (p-value)	Hausman (p-value)
1996-2002	(19)	Pool	-0.23985	0.99665		0.97357	39.5854		
(N=161)		(S.E.)	(0.0414)	(0.0130)			(0.0020)		
	(20)	Fixed		0.99289		0.99328	22.19900		
		(S.E.)		(0.0368)			(0.0000)		
	(21)	Random	-0.23457	0.99441		0.97881		267.6	0.00000
		(S.E.)	(0.0795)	(0.0193)				(0.0000)	(0.9611)
1996-2002	(22)	Pool	-1.49649	1.01119	0.32408	0.98110	45.6081		
(N=161)		(S.E.)	(0.1605)	(0.0111)	(0.0404)		(0.0000)		
	(23)	Fixed		1.02447	1.66771	0.99461	19.01		
		(S.E.)		(0.0300)	(0.4936)		(0.0000)		
	(24)	Random	-1.94981	1.00813	0.44619	0.97952		213.9	7.46000
		(S.E.)	(0.3459)	(0.0170)	(0.0883)			(0.0000)	(0.0240)
1996-2002	(25)	Pool	-0.23406	0.99818		0.96944	34.50		
(N=154)		(S.E.)	(0.0427)	(0.0143)			(0.0020)		
Drop	(26)	Fixed		0.22817		0.99832	125.23		
Russia		(S.E.)		(0.0595)			(0.0000)		
	(27)	Random	0.62869	0.60812		0.97881		263.5	50.34
		(S.E.)	(0.0937)	(0.0260)				(0.0000)	(0.0000)
1996-2002	(28)	Pool	-1.47069	1.00731	0.32006	0.97787	41.2105		
(N=154)		(S.E.)	(0.1651)	(0.0122)	(0.0417)		(0.0000)		
Drop	(29)	Fixed		0.30291	0.83705	0.99866	112.00		
Russia		(S.E.)		(0.0419)	(0.1918)		(0.0000)		
	(30)	Random	-2.02631	0.66663	0.66451	0.97952		215.3	142.51
		(S.E.)	(0.3456)	(0.0232)	(0.0856)			(0.0000)	(0.0007)
1996-2002	(31)	Pool	-0.19998	0.99089		0.96203	23.18		
(N=119)		(S.E.)	(0.0504)	(0.0181)			(0.0000)		
	(32)	Fixed		0.36074		0.99810	139.69		
		(S.E.)		(0.0605)			(0.0000)		
	(33)	Random	0.42750	0.69668		0.97881		270.2	45.27
		(S.E.)	(0.1110)	(0.0342)				(0.0000)	(0.0000)
1996-2002	(34)	Pool	-1.62186	1.01619	0.34724	0.97377	35.5355		
(N=119)		(S.E.)	(0.1991)	(0.0155)	(0.0475)		(0.0000)		
	(35)	Fixed		0.35069	0.74390	0.99833	107.38		
		(S.E.)		(0.0502)	(0.2435)		(0.0000)		
	(36)	Random	-0.84763	0.74763	0.29610	0.97952		231.73	130.5
		(S.E.)	(0.4225)	(0.0302)	(0.1025)			(0.0000)	(0.0007)

Notes: Sample economies included after 1996 are: Argentina, Australia, Brazil, Britain, Canada, Chile, China, Denmark, Hong Kong, Hungary, Japan, Malaysia, Mexico, New Zealand, Poland, Russia, Singapore, South Africa, South Korea, Sweden, Switzerland, Taiwan, and Thailand.

F-value for the pooling model is the test statistic for the joint null hypothesis that $\alpha = 0,\ \beta 1 = 1$.

F-value for the fixed-effects model compares the fixed-effects model versus the null hypothesis of the pooling model. High values of LM favor the fixed-effects model and the random-effects model over the pooling model. High (low) values of the Hausman test statistic favor the fixed-effects model (random- effects model).

Table 4 Exchange Rate Regimes

Country	2001 April	2000 April	1999 March	1998 April	1997 April	1996 April	1995 April	1994 April	1993 April	1992 April	1991 April	1990 April	1989 April	1988 Mar.	1987 Jan.	1986 Sep.
Argentina	1	1	1	1	1	1	1	1	1	1						
Australia	4	4	4	4	4	4	4	4	4	4	4	4	4	4	4	4
Austria				1	1	1	1	1								
Belgium				1	1	1	1	1	1	1	1	1	1	1	1	1
Brazil	3	3	5	2	2	2	2	5	5	5						5
UK	3	3	3	3	3	3	3	3	3	1	1	3	3	3	3	3
Canada	2	2	2	2	2	2	2	2	2	2	2	2	2	2		2
Chile	3	3	3	3	3	3	3	3								
China	1	1	1	1	1	1	1	1	2	3	3					
Czech Rep	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.								
Denmark	1	1	1	2	2	2	2	2	2	2	2	2	2	2	2	
France	1	1	1	1	1	1	1	1	1	1	1	1	1	1	1	2
Germany	1	1	1	4	4	4	4	4	4	4	4	4	4	4	4	4
Hong Kong	1	1	1	1	1	1	1	1	1	1	1	1	1	1		1
Hungary	3	3	3	2	2	2	2	3	3	3	3					
Indonesia	4	4	5	5			2									
Israel		3	3	3	3	3	3									
Ireland									2	2	2	2	2	2	2	2
Italy	1	1	1	1	1	2	2	2	2	2	2	2	2	2	2	
Japan	4	4	4	4	4	4	4	4	4	4	4	4	4	4		4
Malaysia	1	1	1	4	2	2	2	2	2							
Mexico	3	3	3	3	3	3	5	3	1							
Netherlands			1	1	1	1	1	1	1	1	1	1	1	1	1	1
New Zealand	3	3	3	3	3	3	3									
Poland	3	3	3	3	3	3	6	6								
Russia	2	2	5	6	6	5	5	5	5	5	n.a.	n.a.				
Singapore	3	3	3	2	2	2	2	2		2	2	2	2	2		2
South Africa	4	4	4	4	4	4										
Korea	4	4	4	5	2	2	2	2	2	2	2	2	2			
Spain	1	1	1	1	1	1	1	2	2	2	2	2	2	2		2
Sweden	3	3	3	3	3	3	3	3	3	2	2	2	2	2		2
Switzerland	2	2	2	2	2	2	2	2	2							
Taiwan	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.								
Thailand	3	3	3	3	1	1	1	1	1							

Notes: The classification codes of Reinhart and Rogoff (2002) are as follows:

1: No separate legal tender, Pre-announced peg or currency board arrangement, Pre-announced horizontal band that is narrower than or equal to +/-2%, De facto peg.

2: Pre-announced crawling peg, Pre-announced crawling band that is narrower than or equal to +/-2%, De facto crawling peg, De facto crawling band that is narrower than or equal to +/-2%.

3: Pre-announced crawling band that is wider than or equal to +/-2%, De facto crawling band that is narrower than or equal to +/-5%, Moving band that is narrower than or equal to +/-2% (i.e., allows for both appreciation and depreciation over time), Managed floating.

4: Freely floating.

5: Freely falling.

6: Dual market in which parallel market data is missing.

Data source: Reinhart and Rogoff (2002), http://www.puaf.umd.edu/faculty/papers/reinhart/monthly1.dta.

Shaded areas in the table indicate the samples that are included for the analysis of balanced panel data in section 4 (2).

Table 5 Results of Regressions

Dependent Variable = Ln(P(i,t)/P(t)), Sample 1996-2002							
Methods and samples	Model	Methods	α	β 1	β 2	γ	Long Run Elasticity
IV	(37)	Fixed		1.0296	2.9079	0.2950	1.4604
(N=161)		(S.E.)		(0.0371)	(0.5046)	(0.0355)	
MDE	(38)	Fixed		0.9550	1.7026	0.1114	1.0747
(N=161)		(S.E.)		(0.0278)	(0.2487)	(0.0250)	
IV	(39)	Random	-1.47075	0.9752	0.3106	0.0389	1.0147
(N=161)		(S.E.)	(0.6245)	(0.1892)	(0.1537)	(0.1834)	
GLS	(40)	Random	-5.4058	0.9515	1.3084	0.1476	1.1163
(N=161)		(S.E.)	(21.8564)	(0.0426)	(0.3588)	(0.0325)	
IV	(41)	Fixed		0.0341	-0.1741	1.0472	-0.7232
(N=154, drop Russia)		(S.E.)		(0.0805)	(0.3503)	(0.1418)	
MDE	(42)	Fixed		0.1136	0.1764	1.0788	-1.4410
(N=154, drop Russia)		(S.E.)		(0.0351)	(0.0918)	(0.0452)	
IV	(43)	Random	-0.22948	0.2055	0.0467	0.8018	1.0366
(N=154, drop Russia)		(S.E.)	(1.3292)	(0.7633)	(0.2944)	(0.7581)	
GLS	(44)	Random	-0.9633	0.1847	0.3976	0.5200	0.3848
(N=154, drop Russia)		(S.E.)	(1.3907)	(0.0702)	(0.2805)	(0.1411)	
IV	(45)	Fixed		0.0056	0.1234	0.8996	0.0557
(N=119)		(S.E.)		(0.1220)	(0.4537)	(0.1542)	
MDE	(46)	Fixed		0.0659	0.3004	0.9842	4.1735
(N=119)		(S.E.)		(0.0554)	(0.1337)	(0.0616)	
IV	(47)	Random	-0.2969	0.2481	0.0605	0.7602	1.0347
(N=119)		(S.E.)	(1.8982)	(1.0276)	(0.4085)	(1.0121)	
GLS	(48)	Random	-1.4810	0.5368	0.175145	0.5089	1.0930
(N=119)		(S.E.)	(1.4587)	(0.1510)	0.096664	(0.3368)	

Figure 1：Whole sample

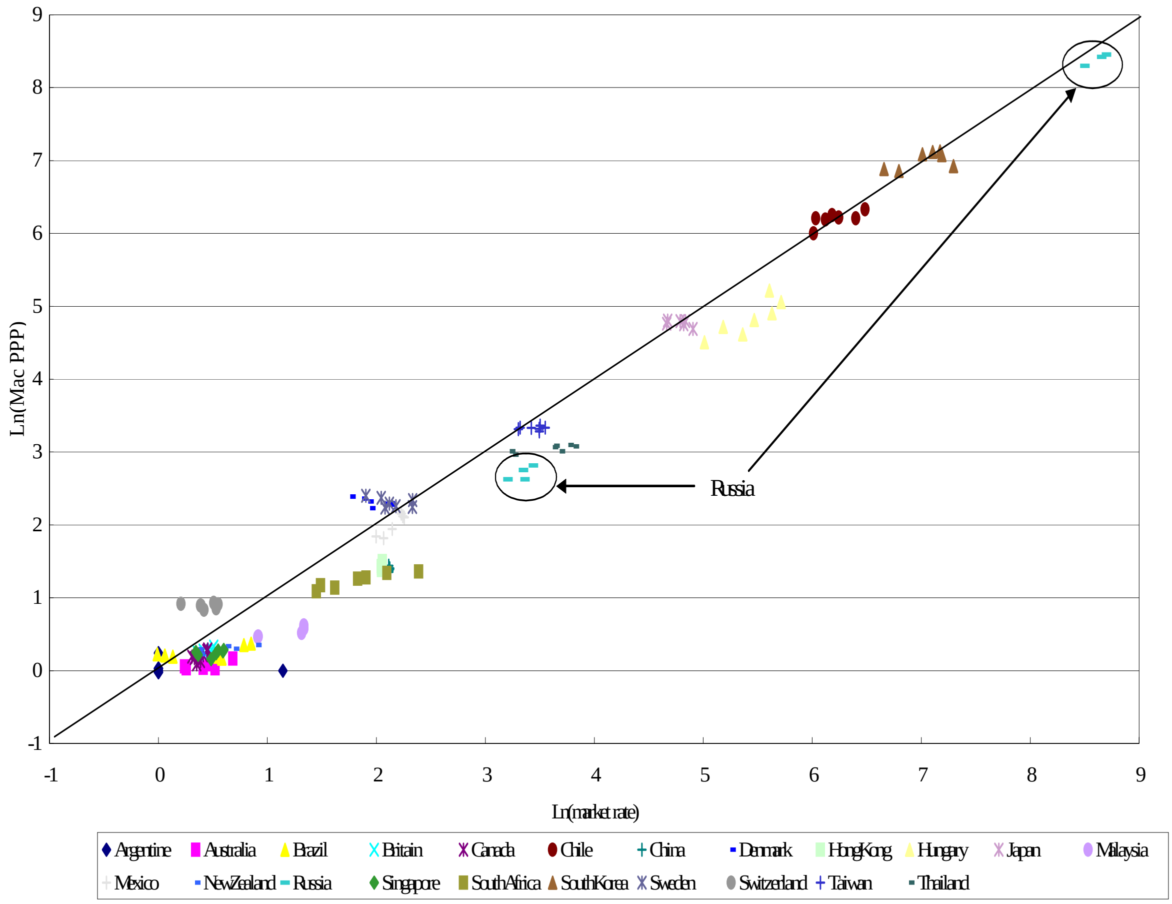

Figure 2: Balanced Panel (Small values only)

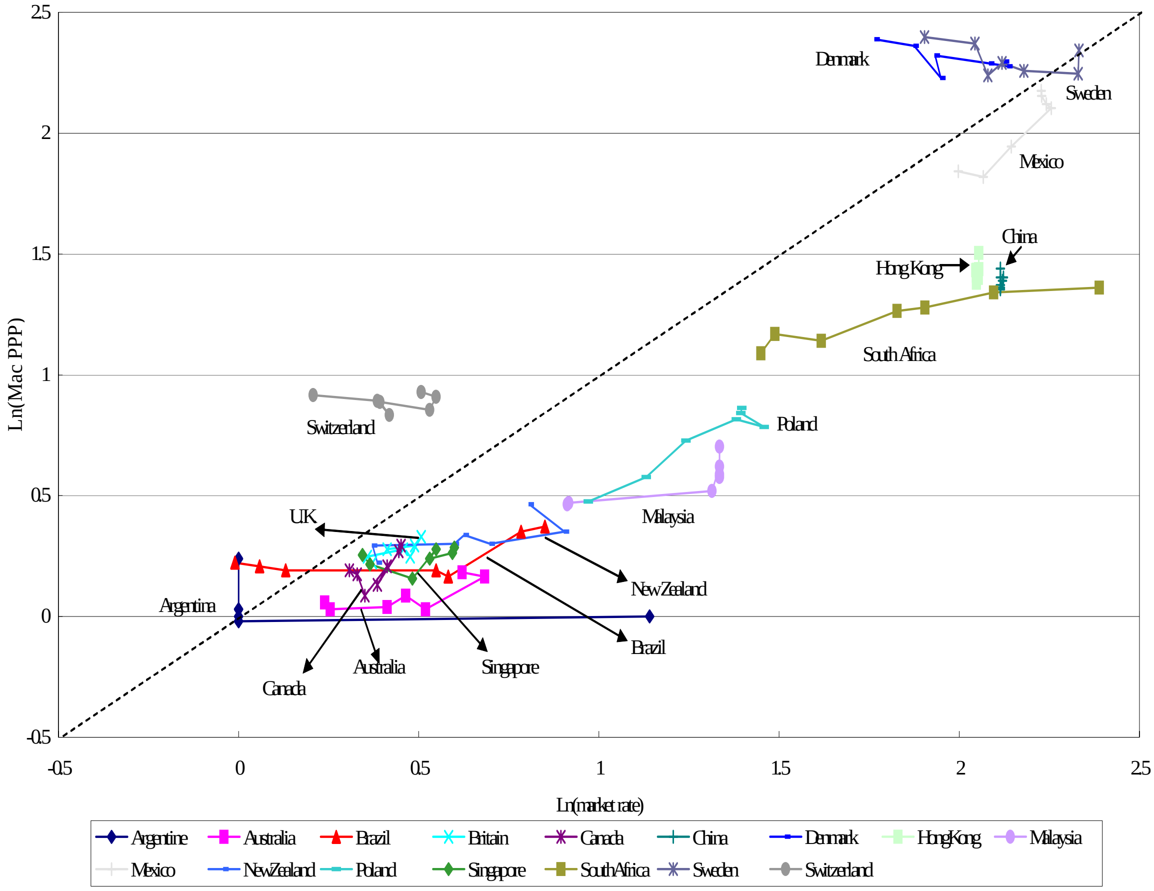

Figure 3: Balanced Panel (Large values only)