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Abstract

We make use of The Economist's Big Mac Index (BMI) to investigate the Law of One Price (LOP) and whether the BMI can be used to predict future exchange rate and price changes. Deviations from Big Mac parity decay quickly, in approximately 1 year. The BMI is a better predictor of relative price changes than of exchange rate changes, and performs best when predicting a depreciation of a currency relative to the US dollar. Convergence to Big Mac parity occurs more rapidly for currencies with some form of exchange rate control than for freely floating exchange rates, which is the opposite of what we obtain using the aggregate CPI.

JEL Classification: F30

Key words: Purchasing power parity; Big Mac index; panel data.

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1 Introduction

In July 2015, the US dollar price of a Big Mac in Switzerland was \$6.82; the price in the US was \$4.79, while the price in Ukraine was \$1.55. There is of course no reason to expect prices to be the same across countries; the Law of One Price is postulated to hold only when homogeneous goods are freely and costlessly traded across countries. Although the Big Mac may be regarded as a homogeneous good, it cannot be regarded as a freely traded good. But whilst the Big Mac is not freely tradeable, many of its ingredients are sourced globally using McDonald's global supply chain. On the other hand, large parts of the cost of producing a Big Mac (wages, rents, utilities) are location-specific (Parsley and Wei, 2007, 2008). Hence, whether there is a tendency towards price convergence across different countries is a relevant question.

The objective of this paper is to evaluate whether the Law of One Price (LOP) holds for the Economist's Big Mac Index (BMI). The Big Mac index has several features that make it a useful means of analysing LOP. First, data is available for many countries over many years. Hence, despite the relatively short time dimension of the data, by making use of the cross-section dimension, we are able to obtain a relatively large sample. This gives our tests more power to reject the null hypothesis of a random walk (non-convergence). Second, the Big Mac is a single, homogeneous good (although Pakko and Pollard (1996) suggest that it is not identical across countries). This is unlike conventional analysis of Purchasing Power Parity (PPP)[1] based on the consumer price index (CPI) or other price indices, which consist of baskets of many different goods, the relative weights of which may be different (and changing) across countries, making cross-country comparisons difficult.

The approach used in this paper is closest to that of Cumby (1997). However, there are several key differences between this paper and Cumby (1997). First, we make use of data since 2004. Whilst maintaining a balanced panel (thus allowing for a battery of panel unit root tests), because more countries have been included in the BMI in recent years, we have the advantage of a much larger dataset than was available to Cumby (1997), with 39 countries and 12 years, as compared with 14 countries and 11 years in Cumby (1997). A second innovation is in the use of dynamic panel data models as in Blundell and Bond (1998), which helps to overcome the bias induced by the lagged dependent variable, and is appropriate since in our dataset the cross-sectional dimension N is larger than the time dimension T. Third, we extend the

[1] The LOP applies to individual goods, whereas PPP applies to the general price level.

basic model to include GDP per capita, to capture the Balassa-Samuelson effect on the real exchange rate. Fourth, following Lutz (2001), we compare the results using the BMI with those using aggregate consumer price indices from the World Bank's World Development Indicators (WDI). Fifth, following Parsley and Wei (1995), we consider nonlinearity in the adjustment towards LOP. Sixth, we conduct some sensitivity analysis to address the issues raised by Froot and Rogoff (1995) regarding the sample used by Cumby (1997).

Since Cumby (1997), there have been many papers on the Big Mac index. This includes Click (1996), Pakko and Pollard (1996, 2003), Ong (1997), Lutz (2001), Lan (2006), Clements and Lan (2010), and Clements et al (2014). Most of the literature has emphasised the time series dimension of the data; for instance, Clements et al (2014) use a sample of 24 countries from 1994 to 2008. Our emphasis is instead on the cross-sectional dimension; by starting the sample in 2004 instead of an earlier starting point, we are able to maximise the number of countries included in the sample. Parsley and Wei (2007, 2008) also make use of Big Mac prices, but they obtain their data from different sources (the Economist Intelligence Unit (EIU) Worldwide Cost of Living Survey and the Mercer Cost of Living survey, respectively). The literature on PPP in general is large, and good surveys include Froot and Rogoff (1995), Rogoff (1996), Taylor and Taylor (2004) and Burstein and Gopinath (2014).

Rogoff (1996) notes that conventional estimates of the rate of decay of deviations from PPP is three to five years. This long half-life, when juxtaposed with the large short-term volatility of real exchange rates, is often referred to as the PPP puzzle (Rogoff, 1996). Proposed solutions to the PPP puzzle include nonlinear models (see Taylor et al, 2001, Taylor and Taylor, 2004), which find in general that the half-life of larger shocks is much shorter (often less than one year) than that of smaller shocks (three to five years). In our results, using either the BMI or CPI, we find half-lives of roughly one year, even when using linear models. Hence the PPP puzzle can be solved without recourse to nonlinear models[2].

The next section presents the data and unit root tests of the data. This is followed by the dynamics of deviations from Big Mac parity and the adjustment of exchange rates and prices in Sections 3 and 4. In Section 5 we conduct sensitivity analysis on our results, and Section 6 concludes. An Appendix gathers together additional results.

[2] This is of course not the same as saying that the true underlying model is linear.

2 Data and Panel unit root tests

There are two main sources of data. Big Mac prices have been published by The Economist since 1986. Initially this was published on an annual basis, usually in April, but since 2012 it has been published bi-annually, in January and July. The number of countries included has increased over time; the original 1986 publication had 15 countries; the July 2015 publication had 57 countries. In this paper, we make use of annual data from 2004 to 2015; where data has been published bi-annually, we use data from the July edition[3]. This yields a balanced panel of 39 countries over 12 years. We combine this with data from the World Bank's World Development Indicators, on per capita GDP at nominal exchange rates, and consumer price indices (CPI). Table 1 presents the countries included in our sample, along with information on the de facto exchange rate regime as of 30 June 2004 (the start of our sample period) as published by the IMF's Classification of Exchange Rate Arrangements and Monetary Policy Frameworks.

Table 1: List of countries in the sample.

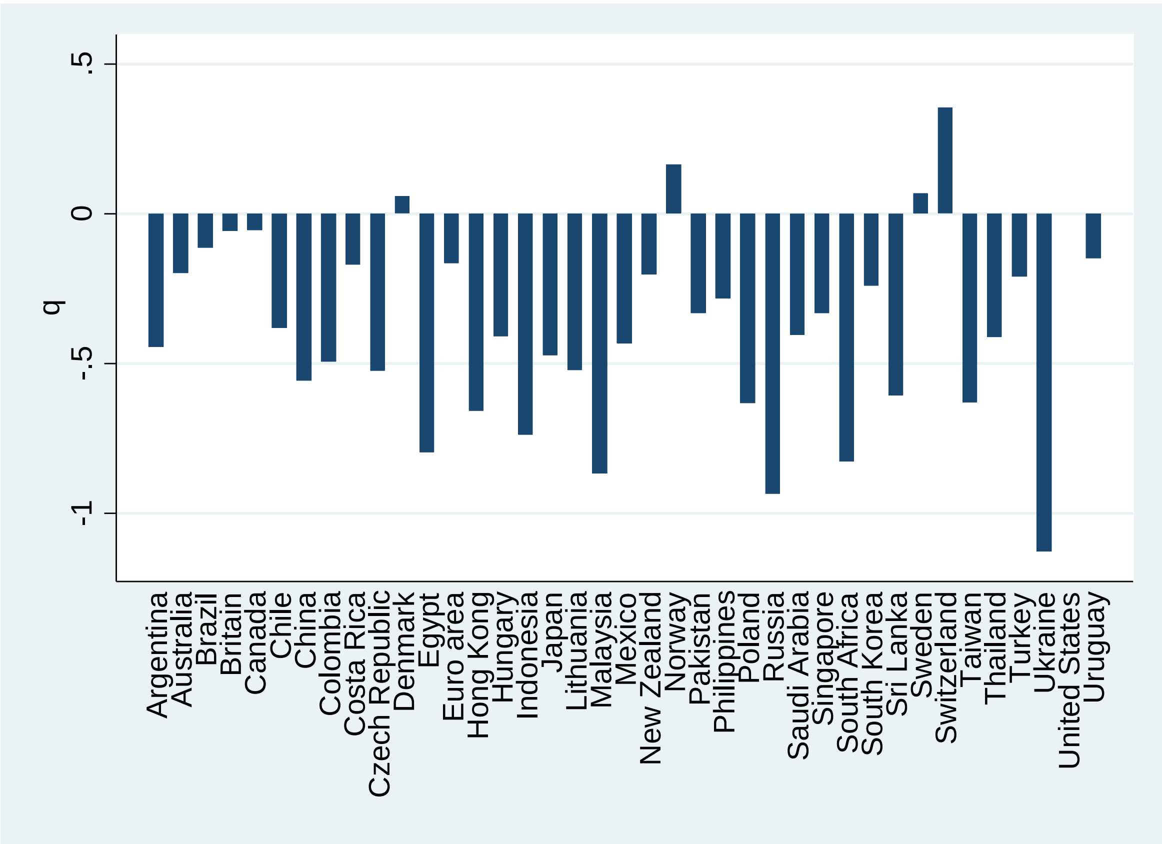
Argentina^	Hong Kong	Saudi Arabia
Australia*	Hungary	Singapore^
Brazil*	Indonesia^	South Africa*
Britain*	Japan*	South Korea*
Canada*	Lithuania	Sri Lanka*
Chile*	Malaysia	Sweden*
China	Mexico*	Switzerland*
Colombia*	New Zealand*	Taiwan^
Costa Rica	Norway*	Thailand^
Czech Republic^	Pakistan^	Turkey*
Denmark	Philippines*	Ukraine
Egypt^	Poland*	United States*
Euro area*	Russia^	Uruguay*

Notes: * indicates a country with an independently floating exchange rate. ^ indicates a country with a managed float. The other countries have some form of peg or crawling peg.

[3] Big Mac prices may vary within countries. For most countries, the Big Mac price is the price in the largest or capital city. For the United States, it is the average of prices in four cities: New York, Atlanta, Chicago and San Francisco. For the Euro area, it is the weighted average of prices in member countries. For China, it is the average of prices in five cities.

We follow the notation used in Cumby (1997). Let E_{it} be the exchange rate between currency i and the US dollar at time t, expressed as units of currency i per US dollar. Let P_{it} be the local currency price of a Big Mac. Define the Big Mac parity exchange rate, $EBMP_{it}$ as the exchange rate that equates the US dollar price of Big Macs in country i with the US price of Big Macs. Thus, $EBMP_{it} = P_{it}/P_{USt}$, and $q_{it} = \ln(EBMP_{it}/E_{it})$ is the approximate percent deviation from Big Mac parity; equivalently, q_{it} is the logarithm of the real Big Mac exchange rate. Figure 1 plots q_{it} for 2015, from which it is clear that the majority of countries in the sample have currencies which are undervalued relative to the US dollar (i.e. $EBMP_{it} < E_{it}$, or $q_{it} < 0$). The exceptions are Denmark, Norway, Sweden and Switzerland. This pattern is not unique to 2015: the average value of q_{it} across our sample of 468 country-year observations is -0.236. For only 118 observations is q_{it} greater than zero.

Figure 1: Values of $q_{it} = \ln(EBMP_{it}/E_{it})$ for 2015.



If LOP holds in the long run, deviations from the parity should be temporary, and hence q_{it} should be a stationary process. To investigate this, we perform two panel unit root tests: (1) the Harris and Tsavalis (1999) (HT) test; and (2) the Im-Pesaran-Shin (2003) (IPS) test. We use these instead of alternatives such as the Levin-Lin-Chu (2002) (LLC) test because our data has a larger cross-section than time dimension; $N > T$. These tests are based on the following regression:

$$\Delta q_{it} = \phi_i q_{it-1} + \boldsymbol{z}'_{it}\gamma_i + \epsilon_{it} \tag{1}$$

The null hypothesis is that $H_0: \phi_i = 0$ (unit root) for all i versus the alternative $H_a: \phi_i < 0$. The term $\boldsymbol{z}_{it}$ can represent panel-specific means, or panel-specific means and a time trend. The HT test requires a balanced panel, and that all panels share the same autoregressive parameter, so that $\phi_i = \phi$ for all i. On the other hand, IPS allows for unbalanced panels, and for the autoregressive parameter to be panel-specific[4]. We perform both tests allowing for a linear time trend, and the cross-sectional averages have been subtracted from the data. The results are presented in Table 2. Both the HT and IPS tests reject the null hypothesis that there is a unit root across all countries, in favour of the alternative that there is at least one country in which q_{it} does not have a unit root.

Table 2: Panel unit root tests of deviations from Big Mac parity, $q_{it} = \ln(EBMP_{it}/E_{it})$.

Test	Test Statistic	p-value
Harris-Tsavalis (HT) $\hat{\phi}$	-4.0214	0.0000
Im-Pesaran-Shin (IPS) $Z_{\tilde{t}-bar}$	-3.5914	0.0002

Notes: $N = 39, T = 11$ for both tests. Both tests include a linear time trend, and the cross-sectional averages have been subtracted from the data.

3 The dynamics of deviations from Big Mac parity

We estimate dynamic panel data models of the following form:

$$q_{it} = \theta_i + \lambda_t + \rho q_{it-1} + u_{it} \tag{2}$$

If there is convergence, $\rho = 1 + \phi$ should be between zero and one. The absolute value of $\rho - 1$ may be interpreted as the annual decay rate for deviations from LOP, and the half-life is $\ln(0.5)/\ln(\rho)$. The dynamic nature of the estimated equation means that the standard fixed effects estimator is downward biased. To overcome this bias, we use the Blundell and Bond (1998) (BB) system GMM estimator, which is consistent when N is large relative to T. We make use of the efficient, two-step,

[4] It appears from the text as if the IPS test is superior to the HT test. This is not necessarily the case: Hlouskova and Wagner (2006) show in a simulation study that, when panels are homogeneous, the HT test outperforms the IPS test in terms of size and power.

forward orthogonal deviations transform version of this estimator, with Windmeijer-corrected cluster-robust standard errors, and include a full set of time dummies to reduce the correlation across countries in the error term[5].

The results of the regression of equation (2) are presented in Table 3. In column (1), we report the estimate of ρ from a conventional two-way fixed effects estimator. The estimated coefficient is 0.44, and is significantly different from zero. This is similar to the results obtained by Cumby (1997), and indicates rapid decay in deviations from LOP, with a half-life of 0.8 years. Columns (2), (3), (4) and (5) report the results of several different specifications of the BB estimator. In column (2), the standard BB estimator with a full set of instruments gives a statistically significant estimate $\hat{\rho} = 0.607$, which implies a half-life of about 1.4 years. Although somewhat slower than the result obtained by Cumby (1997), it is much more rapid than estimates using aggregate price levels which gave rise to the PPP puzzle.

Table 3: The dynamics of deviations from Big Mac parity.

Dep. Var.: q_{it} Estimation method	(1) FE	(2) Full GMM	(3) Short GMM	(4) Collapsed GMM	(5) PCA GMM
q_{it-1}	0.440 (0.094)***	0.607 (0.092)***	0.633 (0.094)***	0.596 (0.165)***	0.643 (0.134)***
R^2	0.44				
$N \times T$	429	429	429	429	429
N	39	39	39	39	39
T	11	11	11	11	11
Number of instruments		76	40	22	30
Arellano-Bond AR(1) test		0.00	0.00	0.00	0.00
Arellano-Bond AR(2) test		0.33	0.28	0.35	0.29
Hansen test p-value		1.00	0.32	0.08	0.22
Number of components					19
PCA R^2					0.91
Kaiser-Meyer-Olkin measure					0.87

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

In column (2), the BB estimate satisfies the Arellano and Bond (1991) test for no second-order correlation in the residuals. However, it also employs a large number of instruments, which, as Roodman (2009a) points out, may weaken the Hansen test of overidentification. This weakening of the Hansen test can be seen from the p-value of

[5] Cumby (1997) uses the Kiviet (1995) corrected fixed effects estimator, which is more appropriate when N is relatively small compared to T. The Kiviet estimator does not allow for other variables to be endogenous or pre-determined. Cumby (1997) uses a Monte Carlo simulation to determine the magnitude of the fixed effects bias in the presence of pre-determined variables.

1.00. Hence in columns (3) to (5), we employ three different methods to reduce the instrument count. In column (3), instead of the full set of GMM instruments, we make use of only the first two lags. This reduces the number of instruments from 76 to 40, and the p-value of the Hansen test from 1.00 to 0.32, but does not significantly change the estimate of ρ. In column (4), we collapse the set of instruments into a single column (see Roodman (2009b) for details). This reduces the number of instruments to 22, but does not significantly change the estimated value of ρ.

Finally, in column (5), we replace the GMM instruments with the principal components of these instruments. All components with eigenvalues of at least 1 are selected, which yields 19 components. This reduces the instrument count to 30, with the estimated value of ρ similar to that obtained using the other GMM methods. This estimator satisfies both the Arellano and Bond (1991) and Hansen tests. The proportion of the variance explained by the principal components is large (0.91), while the Kaiser-Meyer-Olkin measure of sampling adequacy indicates good performance from the principal components. Since the principal-components-based estimator has good properties, for brevity we henceforth report only the results using this version of the BB estimator.

One explanation for differences in prices across countries is the Balassa-Samuelson effect, which states that that rich countries have higher prices than poor countries, especially in nontradeable goods such as hamburgers. The correlation coefficient between q_{it} and the natural log of GDP per capita is 0.68. Therefore, in addition to the basic specification, we also estimate the following equation, where we include GDP per capita measured in current exchange rates as an additional explanatory variable[6]:

$$q_{it} = \theta_i + \lambda_t + \rho q_{it-1} + \eta \ln GDPPC_{it} + u_{it} \tag{3}$$

The prediction is that the price of a Big Mac is higher in a rich country with higher labour costs: $\eta > 0$. Per capita GDP is treated as an exogenous variable in equation (3).

As noted in the Introduction, nonlinear adjustment has been put forward as a potential explanation for the PPP puzzle. We therefore also consider nonlinear patterns in the rate of convergence. Parsley and Wei (1995) add a quadratic term to the estimated equation, giving:

[6] Our use of current year per capita GDP follows much of the literature (Froot and Rogoff, 1995). Using the previous year's per capita GDP as in The Economist (2015), or using per capita GDP relative to the US, gives very similar results to those reported.

$$q_{it} = \theta_i + \lambda_t + \rho_1 q_{it-1} + \rho_2 q_{it-1}^2 + u_{it} \quad (4)$$

However, with $q_{it} \gtrless 0$, a quadratic regression such as equation (4) implies asymmetric adjustment depending on whether the currency is over- or undervalued. For instance, a positive coefficient on the quadratic term would imply that an undervalued currency ($q_{it-1} < 0$) would adjust more quickly to the undervaluation than would be predicted by a linear relationship, but an overvalued currency ($q_{it-1} > 0$) would adjust less quickly than a linear relationship would predict.

One way of overcoming the asymmetric adjustment problem is to use a cubic equation. A cubic equation also has the additional advantage of providing a test of linearity against an ESTAR (exponentially smooth transition autoregressive) model. Following the intuition in Luukkonen et al (1988) and Michael et al (1997), in our first-order autoregressive equation, this test involves estimating the following cubic relationship:

$$q_{it} = \theta_i + \lambda_t + \rho_1 q_{it-1} + \rho_2 q_{it-1}^2 + \rho_3 q_{it-1}^3 + u_{it} \quad (5)$$

And testing the null hypothesis $H_0\colon \rho_2 = \rho_3 = 0$, against the general alternative that H_0 is not valid, using an F-test. We assume that both the quadratic and cubic terms in equation (5) are endogenous, and instrument them in the usual way.

Table 4 presents both the fixed effects and PCA-based system GMM results of equations (3) and (5). Columns (1) and (2) of Table 4 explore the implications of the Balassa-Samuelson model by including per capita GDP[7]. Consistent with the Balassa-Samuelson model, both columns show highly significant positive coefficients; countries with higher per capita GDP also have higher Big Mac prices. Including per capita GDP in the regression reduces the size of the coefficient on q_{it-1} compared to Table 3: the speed of decay is more rapid, at 0.4 years for the fixed effects estimate, and 0.8 years for the GMM estimate.

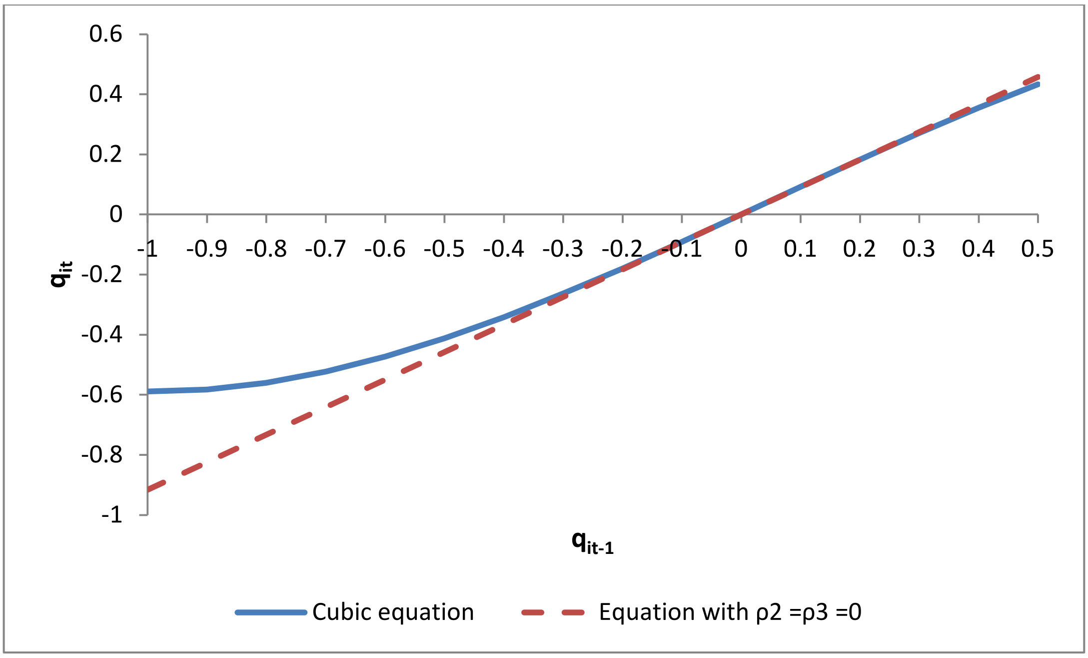
[7] We lose one country – Taiwan – and several other observations due to missing data, and the fact that GDP data are available only up to 2014.

Table 4: Non-linear adjustment, per capita GDP, and deviations from Big Mac parity.

Dep. Var.: q_{it}	(1)	(2)	(3)	(4)	(5)	(6)
Estimation method	FE	PCA GMM	FE	PCA GMM	FE	PCA GMM
q_{it-1}	0.183	0.439	0.597	0.983	0.586	0.916
	(0.068)**	(0.144)***	(0.064)***	(0.126)***	(0.063)***	(0.150)***
$\ln GDPPC$	0.462	0.286				
	(0.060)***	(0.067)***				
q_{it-1}^2					-0.065	0.042
					(0.084)	(0.082)
q_{it-1}^3			-0.179	-0.350	-0.218	-0.285
			(0.063)***	(0.063)***	(0.083)**	(0.099)***
R^2	0.53		0.47		0.47	
$N \times T$	378	378	429	429	429	429
N	38	38	39	39	39	39
T	10	10	11	11	11	11
F-test: $\rho_2 = \rho_3 = 0$			8.18	30.88	4.44	12.24
F-test p-value			0.007	0.000	0.018	0.000
Number of instruments		28		38		50
Arellano-Bond AR(1) test		0.00		0.00		0.00
Arellano-Bond AR(2) test		0.37		0.46		0.42
Hansen test p-value		0.36		0.39		0.93
Number of components		17		26		37
PCA R^2		0.91		0.87		0.89
Kaiser-Meyer-Olkin measure		0.86		0.82		0.75

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

Figure 2: Nonlinear adjustment (equation (5)) versus equation with nonlinear coefficients set to zero.



Notes: A total of 406 observations have values of q_{it-1} between -1 and 0.5. The coefficients used are those from column (6) of Table 4.

The remaining columns of Table 4 consider the implications of nonlinear adjustment. Columns (3) and (4) include only the linear and cubic terms from equation (5), and find a significantly negative coefficient for the cubic term. This implies that currencies with larger deviations from Big Mac parity experience more rapid adjustment than would be predicted by a linear model, and provides evidence in support of the ESTAR model against the linear model[10]. These results are also supportive of an adjustment cost model, in which larger deviations imply faster adjustment. We obtain similar results in columns (5) and (6), which include the linear, quadratic and cubic terms. The quadratic term is never statistically significant, while the cubic term is always so. The F-test of the joint significance of the nonlinear terms is always significant at the 5 percent level or better. Table 4 also reports the Arellano and Bond and Hansen tests, the percentage of variation explained by the principal components, and the Kaiser-Meyer-Olkin measure of sampling adequacy for the GMM models. In all cases the results of these tests are satisfactory. Figure 2 shows the implications of the cubic equation on the adjustment process, relative to an equation in which the nonlinear coefficients are set equal to zero.

[10] For instance, if $q_{it-1} = -0.5$, then the half-life is 3.6 years, whereas if $q_{it-1} = -1.0$, then the half-life is 1.3 years.

Another extension we considered was the role of geographic distance in the adjustment process. The hypothesis is that the closer is a country to the United States, the more quickly deviations from Big Mac parity will decay. However, we find in results reported in the Appendix that distance from the United States plays no statistically significant role in the adjustment process of Big Mac prices.

4 The adjustment of exchange rates and relative local currency prices

In this section, the information content of deviations from Big Mac parity for predicting future exchange rate changes is examined with the following regression:

$$\ln\left(\frac{E_{it}}{E_{it-1}}\right) = \delta_i + \varphi_t + \beta q_{it-1} + \nu_{it} \quad (6)$$

If $\beta \neq 0$, deviations from Big Mac parity provide useful information in predicting exchange rates. In addition to testing the hypothesis that deviations from Big Mac parity give no useful information ($\beta = 0$), it is possible to test whether a currency that is undervalued according to the Big Mac standard ($q_{it-1} < 0$) is likely to appreciate between $t-1$ and t ($\beta > 0$), or whether the appreciation between $t-1$ and t is proportional to the deviation at $t-1$ ($\beta = 1$).

We also investigate whether deviations from Big Mac parity are useful in predicting changes in relative Big Mac prices by estimating the following regression:

$$\ln\left(\frac{EBMP_{it}}{EBMP_{it-1}}\right) = \mu_i + \psi_t + \pi q_{it-1} + \zeta_{it} \quad (7)$$

If a currency appears to be overvalued ($q_{it-1} > 0$), the deviation may be subsequently reduced by a decline in the relative local currency prices of Big Macs ($EBMP$). If $\pi < 0$, current deviations from Big Mac parity are useful in predicting changes in relative local currency prices. In both equations (6) and (7), q_{it-1} is treated as an endogenous variable; making the alternative assumption as in Cumby (1997) that q_{it-1} is pre-determined does not qualitatively change the results reported below[11].

The results are reported in Table 5. In columns (1) and (2), deviations from Big Mac parity do not provide useful information in predicting exchange rates: the coefficients

[11] Equation (6) (respectively (7)) is a restricted version of the equation $\ln E_{it} = \delta_i + \varphi_t + \beta_1 q_{it-1} + \beta_2 \ln E_{it-1} + \nu_{it}$, with the restriction that $\beta_2 = 1$. In the Appendix we present results from the unrestricted versions of equations (6) and (7).

on q_{it-1} are positive but not statistically significant. The magnitude of the coefficients is much smaller than those obtained by Cumby (1997). On the other hand, in columns (3) and (4), deviations from Big Mac parity do provide useful information in predicting changes in relative Big Mac prices. The coefficients are negative as expected, and highly significant. A one percent undervaluation of a currency according to the Big Mac index leads to a 0.43 percent increase in the subsequent relative price of a Big Mac. In all cases, the GMM estimates satisfy the usual specification tests. Hence, unlike the findings of Cumby (1997), we find that deviations from Big Mac parity imply adjustment in relative prices rather than in exchange rates. This is as we may expect; McDonald's is able to influence relative Big Mac prices, but not the exchange rate.

Table 5: Exchange rate and relative price adjustment.

	(1)	(2)	(3)	(4)
Estimation Method	FE	PCA GMM	FE	PCA GMM
Dependent Variable	$\ln(E_{it}/E_{it-1})$		$\ln(EBMP_{it}/EBMP_{it-1})$	
q_{it-1}	0.130	0.109	-0.431	-0.346
	(0.092)	(0.083)	(0.161)**	(0.151)**
R^2	0.05		0.03	
$N \times T$	429	429	429	429
N	39	39	39	39
T	11	11	11	11
Number of instruments		31		31
Arellano-Bond AR(1) test		0.10		0.03
Arellano-Bond AR(2) test		0.56		0.36
Hansen test p-value		0.23		0.54
Number of components		19		19
PCA R^2		0.91		0.91
Kaiser-Meyer-Olkin measure		0.87		0.87

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

We can make use of the estimates from equation (6) to predict exchange rate movements. From Table 5, a currency which is undervalued according to the BMI will appreciate in the following year (albeit not by a statistically significant amount). From Figure 1 and the related discussion above, we conclude that the majority of currencies are undervalued. This suggests that most currencies should appreciate relative to the US dollar. Table 6 reports the performance of the model in qualitative terms, using 2014 data to make a prediction for 2015. In 2014, there is overvaluation ($q_{it} > 0$) for 11 out of the 38 countries (excluding the US). In these countries, the prediction is for the currency to *depreciate* relative to the US dollar, whereas in the other countries the prediction is for the currency to *appreciate*. The rightmost column of Table 6 shows whether the prediction was correct or not. Somewhat disappointingly, the predictions were correct in only 13 of the 38 countries. However, the results are more nuanced than that. In all 11 cases where depreciation was predicted, the prediction was confirmed. Ten of these 11 countries have freely floating exchange rates (the only exception is Denmark, which is pegged to the Euro). The correct prediction of Lithuania's currency appreciation is anomalous, as Lithuania adopted the Euro on 1 January 2015, between the two observations, leading to the large observed change in its exchange rate. Costa Rica is the only country with a crawling peg for which the model yields a correct prediction of an appreciation.

Overall the model works well in predicting a currency depreciation against the US dollar. As noted above in Figure 1, there is overvaluation ($q_{it} > 0$) in 2015 only for Denmark, Norway, Sweden and Switzerland, so we may predict that these countries' currencies will depreciate in 2016. However, for the majority of the countries in the sample, the BMI predicts an appreciation of the currency relative to the US dollar in 2016, which may not be borne out given the continuing weakness of the global economy and the consequent capital flight to the US.

Two further comments are in order concerning the exchange rate predictions above. First, LOP is a long-run theory, so it is no surprise if attempts to predict the following year's exchange rate are relatively unsuccessful. Second, our predictions perform poorly relative to Cumby's (1997) predictions, which correctly predicted the direction of exchange rate movements in nine out of thirteen predictions – a 69 percent success rate as compared with our 34 percent. On closer inspection, all Cumby's currencies had depreciated relative to the US dollar in the prediction period. In nine of the thirteen cases, the Big Mac index correctly predicted the depreciation, but in the remaining four cases, the Big Mac index incorrectly predicted an appreciation. Therefore, similarly to our results, Cumby's (1997) results point to the success of the big Mac index in predicting currency depreciation, but not currency appreciation. The difference is that in Cumby's data, the majority of currencies were predicted to depreciate, whereas in our data, the majority of currencies are predicted to appreciate[12].

[12] Why is there an asymmetry between the success of predictions of appreciation and depreciation? One possible explanation is that, given the current economic weakness in many countries, their currencies will continue to be undervalued. Another possible explanation is that governments and central banks may prefer undervalued currencies in order to stimulate exports. We do not investigate these possible explanations further, but it is an interesting line for future research.

Table 6: Exchange rate predictions.

Country	E_{2014}	q_{2014}	Prediction	E_{2015}	Outcome
Argentina	8.17	-0.62337	Appreciation	9.13	Incorrect
Australia	1.06	0.003985	Depreciation	1.35	Correct
Brazil	2.22	0.199756	Depreciation	3.15	Correct
Britain	0.59	0.026779	Depreciation	0.64	Correct
Canada	1.07	0.090874	Depreciation	1.29	Correct
Chile	564.14	-0.25318	Appreciation	642.45	Incorrect
China	6.2	-0.56454	Appreciation	6.21	Incorrect
Colombia	1847.65	-0.02973	Appreciation	2708.9	Incorrect
Costa Rica	537.3	-0.18091	Appreciation	533.3	Correct
Czech Republic	20.39	-0.3276	Appreciation	24.7	Incorrect
Denmark	5.54	0.070553	Depreciation	6.81	Correct
Egypt	7.15	-0.70561	Appreciation	7.83	Incorrect
Euro area	0.74	0.032529	Depreciation	0.91	Correct
Hong Kong	7.75	-0.68147	Appreciation	7.75	Incorrect
Hungary	228.31	-0.24133	Appreciation	282.88	Incorrect
Indonesia	11505	-0.68033	Appreciation	13344.5	Incorrect
Japan	101.53	-0.27443	Appreciation	123.94	Incorrect
Lithuania	2.56	-0.3177	Appreciation	0.91	Correct
Malaysia	3.17	-0.68843	Appreciation	3.81	Incorrect
Mexico	12.93	-0.38965	Appreciation	15.74	Incorrect
New Zealand	1.15	0.030753	Depreciation	1.51	Correct
Norway	6.19	0.481129	Depreciation	8.14	Correct
Pakistan	98.68	-0.45567	Appreciation	101.7	Incorrect
Philippines	43.2	-0.25836	Appreciation	45.21	Incorrect
Poland	3.07	-0.4705	Appreciation	3.77	Incorrect
Russia	34.84	-0.62962	Appreciation	56.81	Incorrect
Saudi Arabia	3.75	-0.49153	Appreciation	3.75	Incorrect
Singapore	1.24	-0.23351	Appreciation	1.37	Incorrect
South Africa	10.51	-0.72166	Appreciation	12.41	Incorrect
South Korea	1023.75	-0.18006	Appreciation	1143.5	Incorrect
Sri Lanka	130.26	-0.57913	Appreciation	133.85	Incorrect
Sweden	6.84	0.216452	Depreciation	8.52	Correct
Switzerland	0.9	0.353201	Depreciation	0.95	Correct
Taiwan	29.98	-0.59872	Appreciation	31.02	Incorrect
Thailand	31.78	-0.43129	Appreciation	34.09	Incorrect
Turkey	2.09	-0.08071	Appreciation	2.65	Incorrect
Ukraine	11.69	-1.08144	Appreciation	21.95	Incorrect
Uruguay	22.97	0.025625	Depreciation	27.34	Correct

5 Robustness

We perform two robustness checks. First, we divide the sample countries into two groups: one group in which the exchange rates are freely floating, and another group in which they are not. As Froot and Rogoff (1995) note, not all the countries in the sample have freely floating exchange rate regimes, and this may influence the speed of adjustment to deviations from Big Mac parity. From Table 1, of the 39 countries in the sample, 21 have freely floating exchange rate regimes, while the remaining 18 have some degree of restriction to exchange rate movements. We make use of de facto exchange rate regimes at the start of the sample period; although countries do change regime, we maintain the classification of countries throughout our sample.

Table 7: Different exchange rate regimes.

Dep. Var: q_{it}	(1)	(2)	(3)	(4)
Estimation Method	FE	PCA GMM	FE	PCA GMM
Sample	Free float		Not free float	
q_{it-1}	0.588	0.674	0.328	0.456
	(0.048)***	(0.106)***	(0.159)*	(0.386)
R^2	0.59		0.38	
$N \times T$	231	231	198	198
N	21	21	18	18
T	11	11	11	11
Number of instruments		30		30
Arellano-Bond AR(1) test		0.00		0.12
Arellano-Bond AR(2) test		0.35		0.54
Hansen test p-value		1.00		1.00
Number of components		19		19
PCA R^2		0.89		0.93
Kaiser-Meyer-Olkin measure		0.83		0.80

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

The results of regressing equation (2) using the two sub-samples are reported in Table 7. Columns (1) and (2) report the results for the countries with freely floating exchange rates, and columns (3) and (4) the results for countries with non-freely floating exchange rate regimes. The difference in outcomes is interesting. Countries with freely floating exchange rates have large, statistically significant estimates of ρ, with an estimated half-life of deviations of 1.3 years for the fixed effects estimate, and 1.8 years for the GMM estimate. Countries without a freely floating exchange rate have smaller and less significant estimates of ρ, with an estimated half-life of 0.6 years for the fixed effects estimate, and 0.9 years for the GMM estimate. As before, the GMM estimates satisfy the usual specification tests, although now because we have a smaller sample but the same number of instruments, the Hansen test yields a p-value which may indicate the presence of too many instruments (see Roodman 2009a). Overall the results of Table 7 indicate that deviations from Big Mac parity dissipate more quickly in countries where exchange rate movements are restricted. As with some of our previous results, this may be attributed to the fact that

McDonald's is able to directly influence the price of Big Macs, and therefore that adjustment need not necessarily follow what we may expect.

The second robustness check we perform is to compare our results using the Big Mac index with results using the Consumer Price Index from the World Development Indicators[13]. If the theory of PPP is correct, then a price index such as the CPI which includes both tradeable and non-tradeable goods, should converge more quickly than the BMI, which is based on a non-tradeable good. Define the real exchange rate as:

$$RER_{it} = \frac{CPI_{it}}{E_{it}CPI_{USt}} \tag{8}$$

Where E_{it} is the units of domestic currency to buy one US dollar, and CPI_{USt} and CPI_{it} are the US and domestic CPI in time t. We then estimate the analogue of equation (2):

$$\ln RER_{it} = \theta_i + \lambda_t + \rho \ln RER_{it-1} + u_{it} \tag{9}$$

The correlation coefficient between $\ln RER_{it}$ and q_{it} is 0.23. Using a balanced version of the dataset (in which Argentina and Chile were removed), both the HT and IPS panel unit root tests reject the null hypothesis of a unit root in $\ln RER_{it}$ across all panels. Estimates of equation (9) are reported in Table 8. Comparing the full sample results in columns (1) and (2) with the equivalent results in Table 3 shows that deviations in the CPI-based real exchange rate dissipate more quickly than deviations in the BMI, with a half-life of deviations of less than 0.15 and 0.4 years for fixed effects and GMM estimates, respectively. Columns (3) to (6) divide the sample into countries with freely floating exchange rates, and all other countries, as in Table 7. This time, we find that in countries with freely floating exchange rates, deviations from the real exchange rate dissipate very quickly, whereas this is not the case when exchange rates are not freely floating, although even here the half-life of deviations is less than 1.5 years using both estimators. Our results with the CPI under different exchange rate regimes support the idea that exchange rate adjustment plays an important role in restoring equilibrium following shocks. The results also indicate the absence of the PPP puzzle, and in contrast with Lutz (2001), are different from those using the BMI.

[13] The precise variable used is FP.CPI.TOTL, with base year in 2010. We lose 3 countries – Argentina, the Euro area and Taiwan – and a few other observations due to missing data.

Table 8: Dynamics of deviations from CPI-based real exchange rates.

Dep. Var.: $\ln RER_{it}$	(1)	(2)	(3)	(4)	(5)	(6)
Estimation Method	FE	PCA GMM	FE	PCA GMM	FE	PCA GMM
Sample	Full sample		Free float		Not free float	
$\ln RER_{it-1}$	0.009	0.176	0.008	0.184	0.588	0.560
	(0.010)	(0.116)	(0.007)	(0.151)	(0.057)***	(0.324)
R^2	0.45		0.40		0.70	
$N \times T$	355	355	195	195	160	160
N	36	36	20	20	16	16
T	10	10	10	10	10	10
Number of instruments		27		27		27
Arellano-Bond AR(1) test		0.20		0.13		0.07
Arellano-Bond AR(2) test		0.42		0.42		0.27
Hansen test p-value		0.31		0.95		1.00
PCA R^2		0.94		0.91		0.99
Kaiser-Meyer-Olkin measure		0.84		0.82		0.80

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

6 Conclusions

In this paper we investigate whether the Law of One Price (LOP) holds for the Big Mac Index (BMI), and whether the BMI can be used to predict future exchange rate and price changes. Our analysis follows and extends that of Cumby (1997). We make use of more recent data which expands the sample of countries, and different methods that enable us to overcome the endogeneity bias in the regression models. Some of our results are similar to those of Cumby's. We find generally rapid convergence to LOP following any deviation. Deviations from Big Mac parity are more strongly associated with changes in relative prices than in exchange rates. However, the predictive power of the model is limited, and appears to perform well only in predicting future exchange rate depreciation.

In extensions to the basic model, we find that, in line with the Balassa-Samuelson model, deviations from LOP may be explained by per capita GDP, and that including per capita GDP increases the speed of adjustment. Our results with non-linear adjustment indicate that larger deviations from Big Mac parity imply more rapid adjustment. In comparing the results with those obtained using a CPI-based real exchange rate, overall, the BMI adjusts more slowly to deviations than the aggregate CPI. For both BMI and CPI, the half-life of deviations is less than two years, even in linear models; we thus do not find evidence of Rogoff's (1996) PPP puzzle.

For CPI, countries with a freely floating exchange rate regime experience much more rapid adjustment to deviations from PPP than countries in which there is some form of restriction to exchange rate movements. This may be what is expected, and points to the role of the exchange rate in restoring equilibrium in the face of shocks to the system. On the other hand, we get the opposite results for BMI, which may indicate the role of McDonald's in influencing the price of Big Macs. Hence, returning to the question posed in the title of the paper, hamburgers are not harmless: using the BMI instead of the CPI leads to different (but not necessarily wrong) conclusions on PPP.

Perhaps the big takeaway from this paper is that a non-tradeable good exhibits convergence towards LOP, at a rate that is not dissimilar to an aggregate price index. This suggests two directions for future research. On the one hand, future empirical work could use more types of disaggregated goods to investigate in greater detail the implications of tradeability on convergence towards PPP, along lines similar to Parsley and Wei (1996). On the other hand, it suggests that theoretical work needs to take a more nuanced stance towards the tradeable/non-tradeable dichotomy that is often employed.

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Appendix: Additional Results for "Are hamburgers harmless? The Big Mac index in the twenty-first century"

In this section we present some additional results which are only mentioned in passing in the main paper: (1) the effect of distance from the United States on the decay rate of deviations from Big Mac parity; and (2) exchange rate and price adjustment using unconstrained regressions.

A1 Distance from the United States

We extend the basic model in equation (2) to include the natural log of distance to the United States as an explanatory variable, as well as an interaction term between distance and q_{it-1}:

$$q_{it} = \theta_i + \lambda_t + \rho q_{it-1} + \alpha_1 \ln Dist_i + \alpha_2\big((\ln Dist_i) \times q_{it-1}\big) + u_{it} \qquad \text{(A1)}$$

Where the coefficient α_1 captures the effect of distance on deviations from Big Mac parity, while the coefficient α_2 captures the effect of distance on the decay rate of deviations from Big Mac parity. Both distance and the interaction term are assumed to be exogenous.

Table A1: Distance from the United States

Dep. Var: q_{it} Estimation Method	(1) PCA GMM	(2) PCA GMM
q_{it-1}	0.641	0.881
	(0.112)***	(3.125)
$\ln Distance$	0.133	0.159
	(0.134)	(0.132)
$\ln Distance \times q_{it-1}$		-0.042
		(0.347)
$N \times T$	429	429
N	39	39
T	11	11
Number of instruments	30	31
Arellano-Bond AR(1) test	0.00	0.00
Arellano-Bond AR(2) test	0.31	0.38
Hansen test p-value	0.15	0.35
Number of components	19	19
PCA R^2	0.91	0.91
Kaiser-Meyer-Olkin measure	0.87	0.87

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.

The results are presented in Table A1, where we report PCA-based system GMM results[14]. In both specifications reported, distance from the United States has no statistically significant effect on either the deviation from Big Mac parity, or on the speed of adjustment to any deviations. As with other results reported in the text, this indicates that McDonalds' pricing strategy is more sophisticated than a simple diffusion process.

A2 Exchange rates and price adjustment using unconstrained regressions

As noted in footnote 11, equations (6) and (7) may be viewed as restricted versions of the following equations:

$$\ln E_{it} = \delta_i + \varphi_t + \beta_1 q_{it-1} + \beta_2 \ln E_{it-1} + \nu_{it} \qquad \text{(A2)}$$

And

$$\ln EBMP_{it} = \mu_i + \psi_t + \pi_1 q_{it-1} + \pi_2 \ln EBMP_{it-1} + \zeta_{it} \qquad \text{(A3)}$$

With the restrictions that $\beta_2 = 1$ and $\pi_2 = 1$, respectively. In this section, we estimate equations (A2) and (A3) in their unrestricted form, with the lagged dependent variable assumed to be endogenous in each case. These unrestricted equations enable us to investigate the hypothesis of mean reversion in exchange rates and prices. A significant estimate of β_2 between 0 and 1 would indicate substantial mean reversion in exchange rates. If $\beta_2 = 0$, there is complete mean reversion (within the year), and deviations from LOP are purely transitory. If $\beta_2 = 1$, there is no mean reversion, so that the exchange rate follows a random walk.

The results are presented in Table A2, where we report both fixed effects and PCA-based GMM results. The coefficients on the lagged dependent variables in both cases are between 0 and 1, albeit insignificantly different from zero at conventional levels. The magnitude of the coefficients indicate that there is rapid adjustment following shocks to exchange rates and relative Big Mac prices, with a half-life of less than 0.6 years using the GMM estimates.

[14] Distance drops out of the fixed effects model since it is time-invariant, but we are able to recover the coefficients on time-invariant variables under GMM since we estimate the equation both in levels and in orthogonal differences.

Table A2: Unrestricted adjustment of exchange rates and prices.

Dep. Var:	$\ln E_{it}$		$\ln EBMP_{it}$	
	(1)	(2)	(3)	(4)
Estimation Method	FE	PCA GMM	FE	PCA GMM
q_{it-1}	-0.231	-0.376	0.201	0.051
	(0.086)**	(0.295)	(0.096)**	(0.159)
$\ln E_{it-1}$	0.009	0.307		
	(0.026)	(0.227)		
$\ln EBMP_{it-1}$			0.007	0.291
			(0.023)	(0.228)
R^2	0.27		0.09	
$N \times T$	429	429	429	429
N	39	39	39	39
T	11	11	11	11
Number of instruments		40		41
Arellano-Bond AR(1) test		0.04		0.09
Arellano-Bond AR(2) test		0.54		0.18
Hansen test p-value		0.49		0.52
Number of components		28		29
PCA R^2		0.91		0.91
Kaiser-Meyer-Olkin measure		0.80		0.76

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Figures in parentheses are standard errors clustered by country. All regressions reported include a full set of year dummies.