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Pricing-to-market by Brazilian Exporters: a Panel Cointegration Approach*

João Barata Ribeiro Blanco Barroso**

Abstract

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This paper investigates Brazilian exporters pricing behavior, over the long-run, following destination specific exchange rate shocks. The panel cointegration method of Bai, Kao and Ng (2009) is shown to identify the long-run parameter of interest. The method crucially depends on identification and controlling for the common trend in prices to different countries, a trend which is structurally interpreted, like originally proposed by Kneeter (1989), as the exporter's marginal cost. We find evidence of incomplete exchange-rate pass-through in the long-run, which supports the market structure explanations of Krugman (1986), known in the literature as pricing-to-market, over contending short-run sticky-price explanations. The degree of long-run pass-through is also shown to be positively related to technological intensity in the sector, a proxy for low elasticity of substitution of varieties.

Keywords: exchange rate pass-through; panel cointegration, sticky-price, market structure, pricing-to-market.
JEL Classification: F12, F14, F31

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1. Introduction

With sufficiently segmented international markets, exporters may set country specific prices to reflect local demand and competition conditions, a behavior Krugman (1986) called pricing-to-market. The concept is relevant for many empirical puzzles in international economics, such as incomplete exchange rate pass-through to exporter prices and persistent deviations from purchasing power parity. The main advantage over contending explanations, the leading one being sticky prices set by exporters in local currencies, is the optimal character of the pricing rules[1] and the possible accounting for persistent shocks on relative international prices[2].

This paper follows the empirical tradition of identifying microeconomic pricing-to-market behavior by resort to exporter's destination specific mark-up adjustments after exchange rate shocks (Krugman, 1986; Dornbush, 1987; Kneeter, 1989). Advancing previous methods, we look for credible estimates of long run pricing-to-market effects on Brazilian export markets. By definition, the effect cannot be attributed to sticky price and other short run explanations of incomplete exchange-rate pass-through. On the contrary, it would amount to strong evidence in favor of market structure explanations, the more so considering it refers to exporters from a developing country where it is least expected. Additionally, we look for patterns of behavior across different industries, a significant undertaking in face of the large number of imperfect competition models that may explain the results.

Krugman's (1986) original strategy to identify pricing-to-market behavior was trying to control for exogenous trends in prices common across all export destinations in order to recognize divergent trends after a large real exchange rate shock; this was implemented for German export data. Kneeter (1989) developed this basic insight further. The author proposed a panel framework that controls for common trends in prices by the inclusion of a time effect. He also provided structural interpretation for the common trend as the production cost of the marginal unit, such as can be deduced from the optimization problem of a representative firm exporting to many destinations. The

[1] Sticky local currency prices have welfare implications under flexible exchange rate regimes through the addition of a risk premium term to import and export prices [Sutherland (2005)]; by reducing optimal variability, pricing-to-market has a direct bearing on risk premium and welfare.

[2] Atkeson and Burstein (2008) and Ravn (2001) have established the quantitative importance of general equilibrium models with pricing-to-market, as opposed to sticky prices, in reproducing broad features of international relative prices. While these authors focus on segmentation of producer prices on home and foreign markets, we look at segmentation of export price across foreign markets.

model proposed in this paper has very similar interpretation; so, it is worth getting at some details of Kneeter's econometric specification. The author singles out the bilateral exchange rate, measured as the price of exporter currency in each destination country's currency, as the relevant country specific shock, from which to gauge the presence and quantitative magnitude of pricing-to-market behavior. Measuring prices in the exporter's currency, the degree of pricing-to-market can be assessed by the partial effect of the bilateral exchange rate on export price while holding marginal cost fixed. Expressing the variables in logarithms, this coefficient measures the effect of the exchange rate on exporter's mark-up over marginal costs. Also, since foreign prices are just the sum of the exporter price and the exchange rate, one plus the coefficient measures the exchange rate pass-through to foreign consumers due to mark-up adjustments.

An important result from Kneeter (1989) is that U.S. exporters in most industries appear to fully or over pass-through exchange rate movements to foreign consumers. As noted by the author, instead of the usual assumption of foreign demand curve becoming more elastic as price rises, such coefficients would require just the opposite assumption (for mark-ups are inversely related to the elasticity of demand). This sort of result persisted throughout the literature: there is wide dispersion, usually around full pass-through levels and many coefficients have significant but counterintuitive signs. For a recent example, Méjan (2004) found exactly those results for massive data sets on Germany, United States, France, Italy, Japan and United Kingdom exporters, with volume and price data organized in much disaggregated sectors. On a more optimistic tone, Méjan interpreted the wide dispersion as room for structural, microeconomic explanations. In the most recently published paper on the subject, Bugamelli and Tedeschi (2008) report the same pattern of result after many variations of Kneeter's basic specification. Because their estimated coefficients bundle many different products, the dispersion is not as accentuated as in Méjan's paper. Bugamelli and Tedeschi experimented with product classifications in search for some microeconomic rationale for the resulting heterogeneity. They found some evidence of stronger pass-through for products characterized by increasing returns to scale or intense use of science, features often associated with oligopolistic industries. Most of the studies from the literature use annual data to dismiss dynamic considerations. There are attempts at dynamic panel models with quarterly or monthly data, such as Takagi and Yoshida (2001), on Japanese exports. But the inclusion of lagged price as an explanatory variable, with the associated dynamic panel methodology, leads to the same pattern of results.

The main contribution from this paper is to extend Kneeter's panel method to allow for long-run relations between the variables and to actually estimate the long-run parameters. We trust long run relations will provide a clearer picture on pricing-to-market behavior, with coefficients less dispersed, more plausibly signed and easier to connect with microeconomic fundamentals. In addition, the estimation method discriminates between market structure and sticky prices explanations of incomplete pass-through, in the sense that it isolates the former by construction. At a conceptual level, an important advance is to model the common marginal cost trend as a stochastic process on par with the price and exchange rate processes, with the explicit possibility of equilibrium relations among all of the variables. Long-run relations are modeled as cointegration among integrated processes, which is a restricted but useful interpretation. For example, this rules out mean reverting residuals with long memory, as well as overlooks possible inference problems from near unit root series. In a sense, though, the econometric method proposed in this paper is genuinely more general then previous methods used in the literature, since estimates are consistent even if some of the variables are stationary in levels as previous authors have maintained. The unit-root hypothesis is necessary only for a long-run interpretation of the coefficients; even so, the paper tests this null against stationarity with panel techniques that have greater power then single series tests.

The appropriate econometric theory to address long-run issues in a panel framework was only recently developed. Philips and Moon (1996) were the first to propose consistent estimators for panel cointegration vectors with the concomitant development of the asymptotic theory for sequential and simultaneous limits in the two panel dimensions. An important shortcoming was the assumption of independent errors along the cross-section dimension. To overcome this difficulty, Bai, Kao and Ng (2009) developed a "second-generation" framework where common factors, possibly with unit roots, capture the cross-sectional dependence. Since our structural model postulates the existence of cross-section dependence due to the common marginal cost to all export destinations, this seems as an appropriate estimator. Indeed, the behavioral model proposed in this paper has an exact mapping to Bai, Kao and Ng econometric specification. Nevertheless, it should be stressed that, while the common factors are just an econometric device in their model, here we adopt the much stronger structural interpretation of a common marginal cost trend in export prices to different destinations[3]. As already mentioned, the estimator does not require pre-testing the variables for unit-root against the stationarity alternative. However, in order to check if our initial concerns were of any consequence, we also applied Bai and Kao (2002) panel unit root test to the bilateral exchange rate and export price series.

Another contribution from the paper is the use of product classification to uncover possible microeconomic patterns in the estimated pass-through coefficients. Although the approach is similar to Bugamelli and Tedeschi (2008), two important differences have a bearing on the results. First, the authors used unit values throughout, estimating the common effect for a group of products by including an associated dummy variable. In contrast, this paper bundles the products from the start and calculates price indexes on which the whole analysis is conducted. Second, the actual classification scheme is different; the one used here is closely based on technological intensity. As a research principle, the discipline of building the indexes before obtaining the estimates lends more credibility to any pattern eventually found, since one minimizes snooping through many different aggregations. It also potentially ameliorates the measurement error from using unit values as proxies for prices, under the assumption of independent errors. As for the classification system, we have used it in parallel research which actually suggests some connections with structural preference and market structure parameters.

The remainder of the paper is organized as follows. Section 2 describes the data and limitations. Section 3 reports results on panel unit-root tests relevant for interpreting the results. Section 4 is the heart of the paper, explaining the structural model, the econometric model and the estimation results. Section 5 discusses the panel cointegration results in relation to previous estimates and looks for microeconomic patterns. Section 6 concludes.

[3] Hatemi-j and Irandoust (2004) study long-run exchange rate pass-through to Swedish import prices in a cointegrated panel framework; the authors mention “pricing-to-market” although all price data refers to Sweden and have therefore no bearing on segmentation issues; the authors do not admit any structural common factor along the cross-section dimension and use first-generation panel estimators.

2. Data description and limitations

The data sample ranges from the first quarter of 1997 to the third quarter of 2006, which amounts to 39 periods. Depending on the sector, the number of export destination countries can be as little as 29 and as much as 53. On the one hand, the small sample size in the two panel dimensions, time and country, could raise problems for the asymptotic inferences. On the other hand, the estimator used here was shown to have good finite sample properties in simulation experiments reported by Bai, Kao and Ng (2006). Indeed, the mean bias and standard deviation of the estimates keep their good asymptotic properties in samples as small as 20 in both panel dimensions. More importantly, the estimator has much better properties in small samples than the alternatives[4].

As shown in Table 1, Brazilian exports of manufactures were classified in thirteen sectors according to technological intensity. There are many reasons to adopt some level of aggregation. First, this reduces data to manageable proportions. Second, price indexes average out measurement error in the price data. Third, it permits to investigate a possible relationship between pass-through and technological intensity, throwing some light on the microeconomic structure driving the results. Finally, Brazilian official institutions often use this classification scheme, which was developed by OCDE in 1995 (Classification of High-technology Products and Industries), thus facilitating research communication. Some sectors from the original classification, namely aviation, ship building and oil, have been excluded due to insufficient number of export destination countries. The sector share in the table refers to the value exported in the third quarter of 2006 relative to total manufacturing export.

Table 1 also reports the number of countries in each sector. The criteria for including a country in a particular sector were positive export to this country in all periods and for more than half of the products from the sector. This selection rule ensures a balanced panel in each sector, as well as high quality export price indexes. On the other hand, the rule could lead to selection bias. However, the most likely reason for low trade volume is trade barrier, and the pricing behavior exporters would have in case of no barrier should be independent of the actual level of the barrier. Additionally,

[4] See the Monte Carlo section in Bai, Kai and Ng (2006), with particular attention to table 1 and 2 and the column referring to the fully modified continuously updated estimator. Another justification for relying in such small samples is the freshness of the method and the possible implications for an important field of applied research.

excluded countries get none or just a small share of Brazilian exports, and it seems justified to give much less weight to these countries in the estimation of the pooled coefficient. The last section of the paper discusses the selection bias issue further.

Table 1. Industrial sectors by technological intensity: share and number of countries

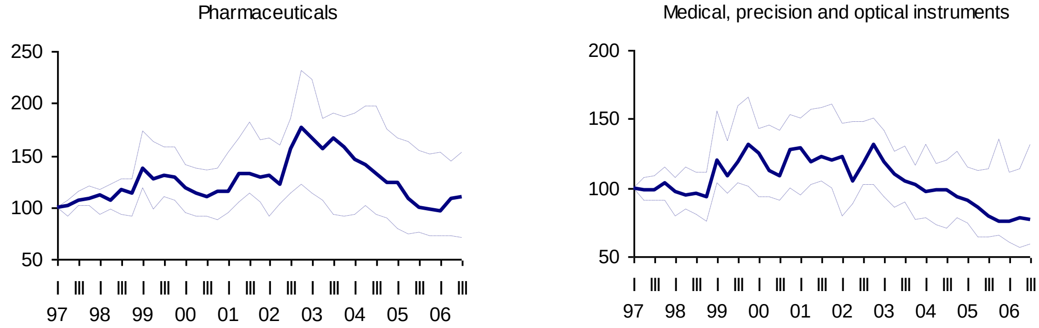
Technology	Industrial Sector	Share (%)	Countries
High	Pharmaceuticals	0,7	29
	Medical, precision and optical instruments	0,5	32
Medium-high	Eletrical machinery and apparatus	2,3	39
	Machinery and equipment	7,1	51
	Chemicals excluding pharmaceuticals	5,8	43
	Motor vehicles, trailers and semi-trailers	11,8	41
Medium-low	Rubber and plastics products	1,8	53
	Non-metallic mineral products	1,8	53
	Basic metals and fabricated metal products	12,8	34
Low	Food products, beverages and tobacco	17,9	46
	Wood, pulp, paper, paper products	5,4	48
	Other Manufacturing	1,0	33
	Textiles, leather and footwear	4,8	52
		73,7	29-53

Export price indexes were calculated with unit values at the finest level of the Harmonized System, an international standard for commodity classification. The Fisher index formula was used after a preliminary trimming of too extreme variation, very unlikely to reflect price developments. For each country, in each sector, an export price series was constructed. Unit values are poor measures of price. Still, they are readily available and very much used in empirical studies on international prices. The inclusion of country specific effects should capture any measurement error that survives aggregation. As already mentioned, one reason to use aggregate price indexes is to average out the likely measurement errors in prices. On the other hand, the more aggregated sector classification we use, the less reasonable our structural interpretation of the data. Relative to other studies with a similar panel approach, our choice in this trade-off involves more aggregation.

Figure 1. Export price indexes by industrial sector; high and medium high technology

Median in bold-line; other quartiles in dashed-lines; 1997:I = 100; Monetary Unit = R$

High Technology



Medium-high Technology

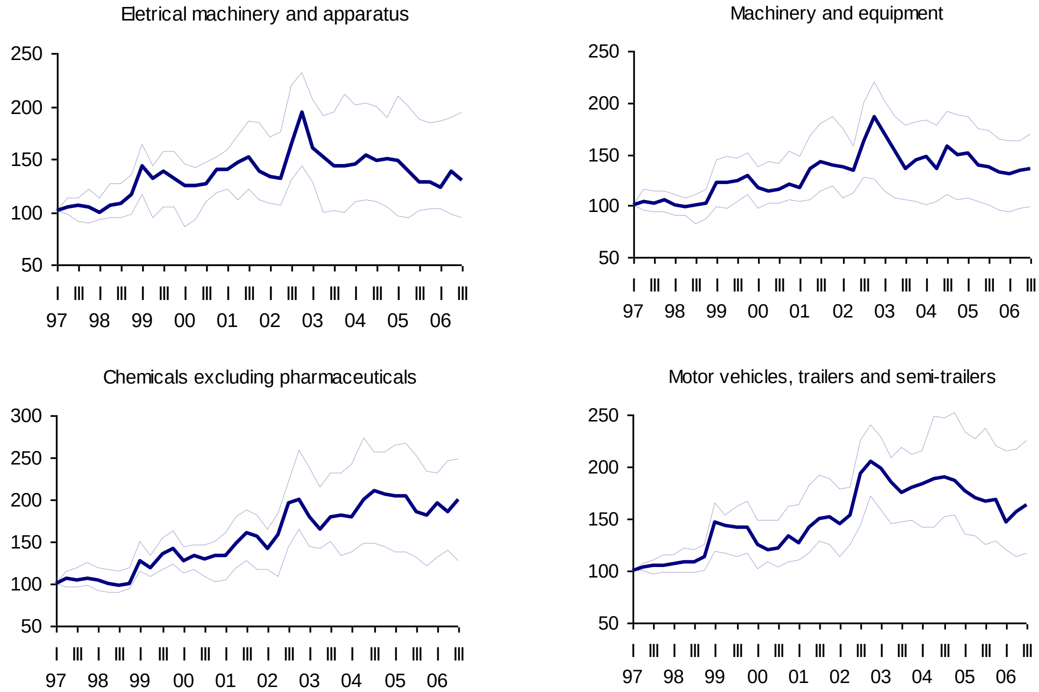
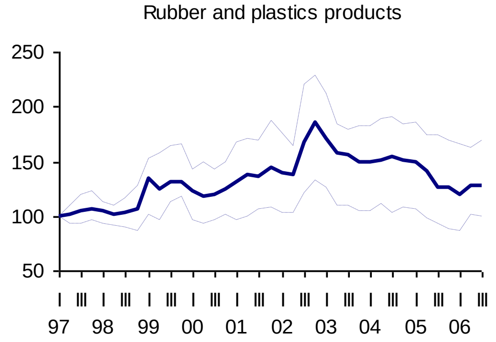


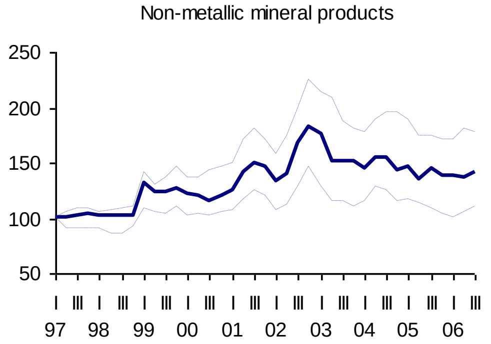
Figure 1 and 2 summarize the export price data for each sector. The median is a rough estimate of a common trend component, while the other quartiles capture how closely country data follow this trend. As can be seen by the widening distance of the quartile-band in most sectors, the common trend appears to have a weak influence in the price data. This will be more formally studied in the next section.

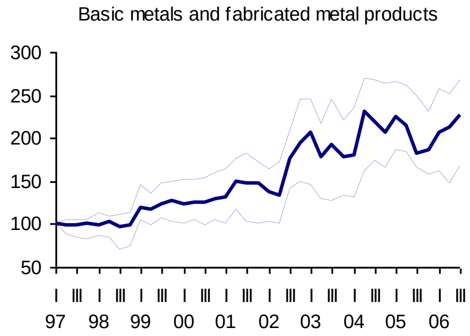
Figure 2. Export price indexes by industrial sector; medium-low and low technology

Median in bold-line; other quartiles in dashed-lines; 1997:I = 100; Monetary Unit = R$

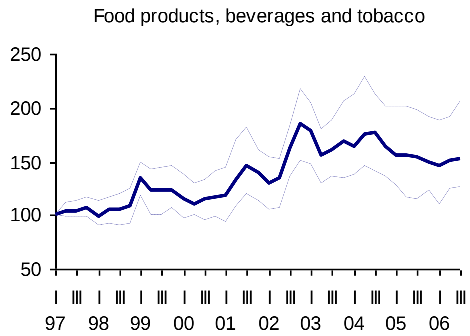
Medium-low Technology

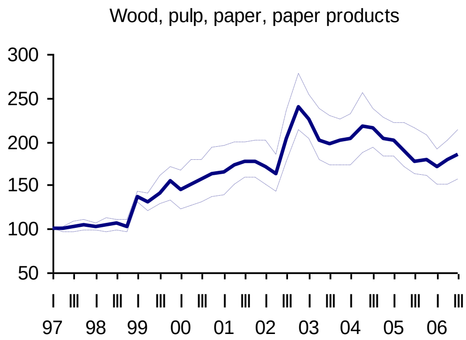


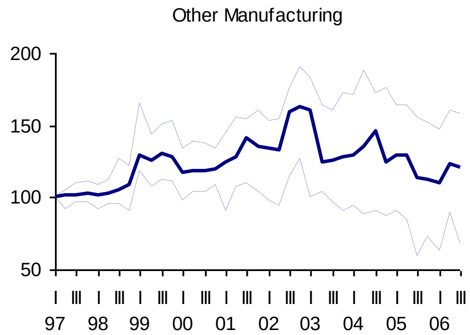




Low Technology







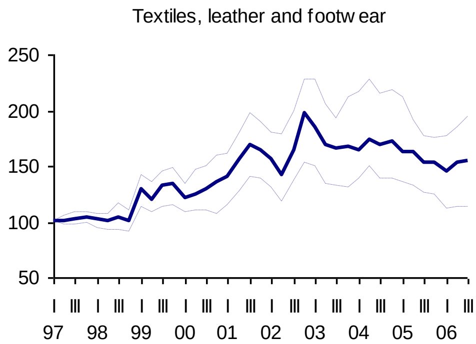
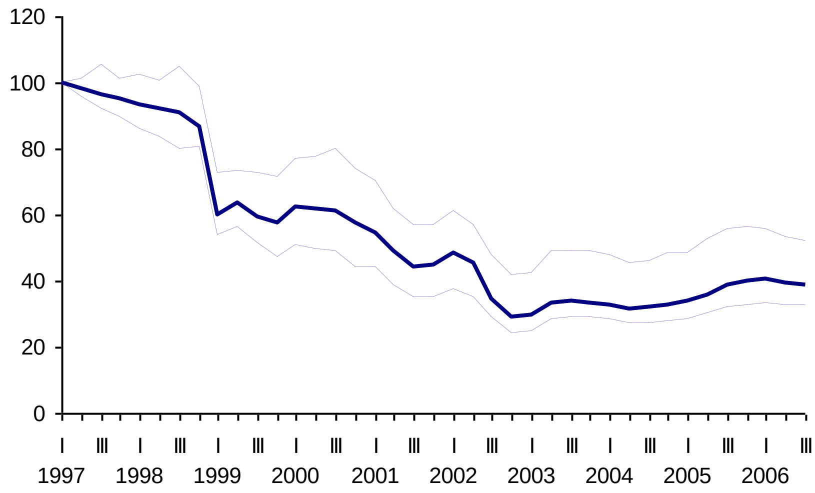


Figure 3. Real bilateral exchange rate

Median in bold-line; other quartiles in dashed-lines; 1997:I = 100



Bilateral exchange rate series were corrected for consumer price inflation in each export destination country. Both the nominal exchange rate and consumer price data were obtained from the International Financial Statistics (IFS) database. After the correction, we obtain a „real' exchange rate measure, from the point of view of consumers in each export destination country. Indeed, we will often refer to the corrected exchange rate as the bilateral real exchange rate. One could argue that sector specific price inflation in each foreign country should be used in place of consumer price inflation, but this kind of data is not easily obtainable. Figure 3 summarizes exchange rate data. Compared to the price series, the common trend captured by the median seems to be much more important for exchange rates, with all the country series following it very closely.

3. Results on panel unit-root

Before estimating the parameter of interest for each industry, we test all the series for unit-root using a panel technique developed by Bai and Kao (2004). The procedure is different from the usual time-series ones in at least two respects; first, it can spot low frequency movements that would otherwise go undetected, as long as they are shared by the series in the panel; second, it can pool the unit-root tests, making it harder to incorrectly accept the unit-root null for a group of time-series. The first feature, which is unique to Bai and Kao approach, will be particularly important for correctly classifying the export price index series. It is important to highlight that none of this pre-testing is necessary for the estimation of the pricing-to-market coefficient; the estimator used in the next section would be consistent even if some of the series were stationary (but it would converge more slowly). Still, pre-testing is crucial for interpreting correctly the estimated parameters.

The test decomposes each series from a panel into a common factor and an idiosyncratic component. The estimation procedure makes no assumption about integration order of the components, thus allowing testing each component separately. Common factors serve as an econometric device to capture cross-section dependence, thus overcoming the most serious deficiency of first-generation panel tests for unit-root. In contrast to the model in the next section where the structural dependency between variables is explicitly addressed, in this section we model each variable in separation from the other.

Formally, each time-series from a given panel of series is decomposed as:

$$y_{it} = d_{it} + \lambda_i' F_t + \nu_{it} \qquad (1)$$

where $i=1..N$ indexes the panel unit, $t=1..T$ indexes the time periods, d_{it} is a determinist trend, F_t is a vector of common factors, λ_i a vector of factor loadings, ν_{it} is the error component and y_{it} is the variable of interest. The common factors are estimated as principal components from the differenced series summed up to each period. The number of factors to be included in the model for each panel was selected by information criteria as suggested by Bai and Ng (2002).

For the export price series, each sector was treated as a separate panel data set, and each of these panels was modeled as in equation (1), with *i* indexing the export destination country. This is a reasonable setup given that the common factors influencing the price to each destination country is most likely sector specific; we have just imposed this by assumption. Admitting at most three factors, a single factor was selected in all panels.

For the exchange rate data, the common factors are mostly associated with macroeconomic conditions unrelated to the sectors. Thus, all the export destination countries were pooled in a single panel, as in equation (1), with *i* indexing the country. Again, admitting at most three factors, a single factor was selected for the exchange rate.

Since there are many panels, each with several time-series, we report only summary information on the testing results. Qualitative results are not sensitive to the specification of the deterministic component; but we report results under different assumptions.

Table 2. PANIC test for unit root; one factor; constant and time trend included *39 time periods*

Export price by technology and sector		Countries	Explained by Factor >25%	Explained by Factor >40%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
High	Pharmaceuticals	29	34%	10%	10%	-2,86	0	28%	>
	Medical, precision and optical instruments	32	16%	3%	18%	-1,63	1	34%	>
Medium-high	Eletrical machinery and apparatus	39	8%	3%	20%	-4,59	1	26%	>
	Machinery and equipment	51	4%	2%	27%	-3,28	0	37%	>
	Chemicals excluding pharmaceuticals	43	53%	28%	23%	-3,00	0	40%	>
	Motor vehicles, trailers and semi-trailers	41	24%	5%	20%	-2,02	2	37%	>
Medium-low	Rubber and plastics products	53	4%	2%	15%	-4,52	1	25%	>
	Non-metallic mineral products	53	43%	17%	19%	-3,43	0	30%	>
	Basic metals and fabricated metal products	34	18%	6%	35%	-2,94	3	47%	>
Low	Food products, beverages and tobacco	46	13%	4%	15%	-2,61	2	24%	>
	Wood, pulp, paper, paper products	48	73%	46%	8%	-2,14	0	19%	=
	Other Manufacturing	33	6%	6%	27%	-2,42	1	39%	=
	Textiles, leather and footwear	52	50%	29%	17%	-2,64	0	31%	>

Bilateral exchange rate corrected for consumer price	n	Explained by Factor >75%	Explained by Factor >90%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
	68	76%	44%	0%	-1,07	0	11%	=

Notes: (i) "Explained by Factor" is the fraction of n countries where the common factor explains at least 25, 40, 75 or 90 percent of the variation. (ii) "%Rej" is the fraction of the original or error series for which the unit-root hypothesis is rejected. (iii) The Dickey-Fuller t-test statistic and lag order are shown only for the common factor, and rejection is indicated be an underline. (iv) A greater

Table 2 reports the results when country specific intercept and time-trend are both included. The column entitled "Explained by Factor" reports the fraction of countries for which the common factor explains at least 25 or 40 percent of the variation in the original series, for the case of export price indexes, and at least 75 or 90 percent of the variation, for the case of the exchange rate series. These bands were selected to highlight where most of the action is occurring. Confirming the informal impression from the last section, the common factor accounts for very little of the variation in the export price series in all sectors, but is the main driving force of exchange rates.

The last main columns of Table 2 refer to Augmented Dickey-Fuller t-tests applied respectively to the original series, to the estimated common factor and to the estimated error. The test for the error series does not include a constant and neither a time trend (the critical value is -1.94). Both deterministic components are included in the tests for the original and the factor series (the critical value is -3.41). The number of lags was selected by sequential F-tests, with a liberal 50% significance level, admitting at most three lags.

In the case of the original series and the error series, only the fraction of rejections of the unit root hypothesis is reported. Thus, the percentage indicates the fraction of countries with stationary series. In the case of the common factor series, the table reports the test statistic and the number of selected lags; rejection is indicated in boldface.

Finally, the last column is the result of a pooled test for the unit-root null on the residual series (tests can be pooled on the residuals because all common factors where already extracted). Only series regarded as non-stationary according to the single-series test where included in the pool. Therefore, the idea is to determine if at least one of this series is actually stationary. Accordingly, the "greater than" sign indicates when the fraction of stationary series is likely larger than indicated in the previous column.

Analyzing the lower part of Table 2, we conclude there is strong evidence of unit-root for all the exchange rate series. In fact, the panel testing procedure reinforces results obtained for the original time series. The main source of non-stationarity is the common factor, although most of the idiosyncratic components also display a unit-root.

Results are more interesting for the export price indexes, as displayed in the upper part of Table 2. One of the motivations for Bai and Kao (2004) was the possibility of a very small non-stationary component, which would be hard to capture by traditional tests. This appears to be exactly the case for the export price series. While the common factor explains little of the variation, they have a unit-root in most sectors. In comparison, the idiosyncratic components are often stationary and explain a lot of the variation in the original series, making it hard to detect a unit-root in the series. Indeed, as can be seen from the fifth column from Table 1, single-series tests often reject a unit-root, completely missing the presence of the low frequency common component. In the three sectors where the common factor appears to be stationary, the idiosyncratic components have unit-roots for 70% or more of the destination countries.

Table 3. PANIC test for unit root; only constant included *39 time periods*

Export price by technology and sector		Countries	Explained by Factor >25%	Explained by Factor >40%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
High	Pharmaceuticals	29	34%	10%	0%	-2,87	0	3%	>
	Medical, precision and optical instruments	32	16%	3%	13%	-1,18	1	16%	>
Medium-high	Eletrical machinery and apparatus	39	8%	3%	15%	-3,98	1	15%	>
	Machinery and equipment	51	4%	2%	22%	-1,39	3	10%	>
	Chemicals excluding pharmaceuticals	43	53%	28%	9%	-1,87	0	14%	>
	Motor vehicles, trailers and semi-trailers	41	24%	5%	17%	-3,55	0	7%	>
Medium-low	Rubber and plastics products	53	4%	2%	8%	-4,52	1	8%	>
	Non-metallic mineral products	53	47%	21%	19%	-2,24	2	8%	>
	Basic metals and fabricated metal products	34	35%	6%	6%	-1,66	1	9%	>
Low	Food products, beverages and tobacco	46	13%	4%	11%	-2,19	2	9%	>
	Wood, pulp, paper, paper products	48	73%	50%	2%	-1,32	0	6%	>
	Other Manufacturing	33	6%	6%	18%	-1,88	1	21%	=
	Textiles, leather and footwear	52	52%	29%	2%	-2,63	0	6%	>

Bilateral exchange rate corrected for consumer price	n	Explained by Factor >75%	Explained by Factor >90%	Series %Rej	Factor t	Factor p	Error %Rej	Error Pool
	68	94%	73%	4%	-1,96	0	10%	=

Notes: (i) "Explained by Factor" is the fraction of n countries where the common factor explains at least 25, 40, 75 or 90 percent of the variation. (ii) "%Rej" is the fraction of the original or error series for which the unit-root hypothesis is rejected. (iii) The Dickey-Fuller t-test statistic and lag order are shown only for the common factor, and rejection is indicated be an underline. (iv) A greater

Table 3 shows results without the time trend. The general picture is much like before, with the exception of a larger proportion of non-stationary idiosyncratic components. There is also disagreement on the integration order of common factors for two of the industrial sectors. But common factors still account for very little of the variation in the export price series, the opposite being the case for the exchange rate

4. Results on panel cointegration

We present coefficient estimates for the partial effect of the bilateral exchange rate on export price controlling for marginal cost. The most direct way of attributing meaning to this coefficient is through the first order condition for the profit maximization problem of a representative domestic firm from a given industry exporting to several destinations. The bilateral exchange rate enters the first-order condition of the respective destination, and may be interpreted as a demand or cost shift. Comparative statics imply mark-up adjustments specific to each market. Equilibrium effects may be incorporated implicitly in residual demand curves or explicitly in additional first order conditions. In this paper, we consider an approximation of the bilateral exchange rate comparative statics effects.

There are many ways to model this formally (see the appendix). The essential point, though, is variable demand elasticity along the demand curves. Kneeter (1989) presents a simple model for an arbitrary residual demand function. Dornbush (1987) explores the residual demands that emerge from different market structures. Atkeson and Burstein (2008) take demand elasticity as a function of market share and preferences for variety. The structural interpretation is essentially static, but the inclusion of a time trend will hopefully capture changes in equilibrium market structure. The next section discusses structural models in greater detail to anchor theoretical expectations on microeconomic fundamentals.

The econometric model for the panel data set for an arbitrary sector is,

$$p_{it} = d_{it} + \beta e_{it} + \lambda_i MgC_t + u_{it} \qquad (2)$$

where $i=1..N$ indexes the country, $t=1..T$ indexes the time periods, d_{it} is a determinist term, p_{it} is the export price in domestic currency, e_{it} is the real bilateral exchange rate (real price of domestic currency in country i), MgC_t is the common marginal cost trend and u_{it} is an error term. Except for the error term, all random variables are assumed integrated of order one; this is in conformity with results from the last section. The error term u_{it} may be at most weakly dependent on the time and the cross-section dimension. In the appendix, we show how this equation may be deduced from first principles.

The coefficient β measures the partial effect of the bilateral real exchange rate, and is the parameter we are interested at. This coefficient is pooled across destination countries. As a consequence, in case there is significant heterogeneity of firm behavior across markets, the parameter actually represents an average effect - as noted by Philips and Moon (1996). The coefficient λ_i measures the partial effect of the marginal cost. It is country specific, which allows for different compositions of export bundles to different countries, or any other reason for some degree of freedom in the assessment of marginal cost relevant for a destination country.

The model follows very closely Bai, Kao and Ng (2009), with the marginal cost trend in equation (2) standing for a common factor in their terminology. Given our desired structural interpretation, we did not use information criterion to set the number of common factors, just fixing it at one as in equation (2). Except for low-technology goods, we argue below results are not very sensitive to this assumption. As for the deterministic term, we have tried the model with and without a deterministic trend. The qualitative results are reasonably close to each other, but with the trend better adhering to the data. The estimation method once again uses principal components to extract the common factor implicit in the export price indexes. Since this allows one to recover the marginal cost trend only up to a linear transformation, there is no expected sign pattern on λ_i's coefficients. For this reason, we will not emphasize these estimates, rather concentrating on the exchange rate effect.

As a matter of comparison, we also estimated two alternative models. The first is Kneeter's specification, for which variables enter in first-difference and the marginal cost trend enter as a fixed time effect (differencing is necessary, given our results on unit-roots from the previous section). The second is Philips and Moon (1996) fully modified estimator, which amounts to model (2) with λ_i's restricted to be zero; that is, disregarding any cross-sectional dependency.

The first results are presented on Table 4. They were obtained under the assumption of both a constant and a time trend present in the country specific determinist term. To obtain the degree of exchange rate pass-through to foreign consumers, just add one to the exchange rate coefficients on the table.Like the previous empirical literature, Kneeter's specification leads to very high levels of pass-through and to a few counterintuitive signs. The Philips and Moon specification appears to be more sensible, but it does not account for cross-sectional dependency (which is very likely given the results from the previous section), and it does not attempt to control for marginal cost, which is necessary for an economic interpretation.

The estimator from Bai, Kao and Ng addresses both issues simultaneously. Their fully modified continuously updated estimator, which has a normal asymptotic distribution, is reported on the fifth column, with robust standard errors bellow it. All coefficients are negative at any reasonable significance level. The degree of pass-through implied by the coefficients is much more plausible then the one implied by previous methods. Comparing the fifth and the third column, we see that cross-sectional dependence is an important issue for most of the sectors. Now comparing the fifth with the fourth column, we notice that endogeneity of regressors, probably of the common marginal cost trend, may be an issue in some industrial sectors. Given the superior performance of the fully modified continuously updated estimator in the Monte Carlo experiments conducted by Bai, Kao and Ng (2009), we take this as our final estimator, and all the other summary measures in Table 4 take this reference point.

Table 4. Panel cointegration results; one factor; constant and time-trend included *39 time periods*

			Alternative estimates			Final estimate (β)			
		Countries	ols	fm-ols	cup	fm-cup	fit	e/cmg	%Rej
High	Pharmaceuticals	29	-0.19	-0.53	-0.44	-0.4 (0.038)	54%	1.31	69%
	Medical, precision and optical instruments	32	-0.1	-0.49	-0.17	-0.26 (0.05)	45%	0.26	63%
Medium-high	Eletrical machinery and apparatus	39	-0.1	-0.47	-0.24	-0.23 (0.046)	45%	0.4	54%
	Machinery and equipment	51	-0.18	-0.44	-0.4	-0.38 (0.024)	31%	2.72	61%
	Chemicals excluding pharmaceuticals	43	-0.05	-0.35	-0.45	-0.46 (0.026)	38%	7.23	60%
	Motor vehicles, trailers and semi-trailers	41	-0.08	-0.39	-0.53	-0.53 (0.026)	55%	3.69	68%
Medium-low	Rubber and plastics products	53	-0.08	-0.38	-0.52	-0.46 (0.03)	52%	2.21	55%
	Non-metallic mineral products	53	-0.11	-0.4	-0.42	-0.4 (0.022)	49%	3.1	66%
	Basic metals and fabricated metal products	34	0.03	-0.25	-0.41	-0.53 (0.039)	45%	2.92	56%
Low	Food products, beverages and tobacco	46	-0.1	-0.28	-0.5	-0.46 (0.024)	44%	2.55	61%
	Wood, pulp, paper, paper products	48	-0.07	-0.58	-0.56	-0.59 (0.024)	80%	3.07	56%
	Other Manufacturing	33	-0.14	-0.41	-0.41	-0.38 (0.033)	36%	4.77	64%
	Textiles, leather and footwear	52	0.02	-0.45	-0.45	-0.43 (0.02)	47%	2.81	69%

Notes: (i) "ols" is the estimator from Kneeter. (ii) "fm-ols" is the estimator from Philips and Moon. (iii) "cup" is the continously updated estimator from Bai, Kai and Ng. (iv) "cup-fm" is the fully modified version of the last estimator, with standard error in parentheses. (v) The fit is the median explained variance among destination countries. (vi) The "e/cmg" is the median ratio of variances of the exchange rate effect and the marginal cost effect. (vii) "%Rej" is the fraction of rejections of the unit-root null aplied to model residuals of each country in the sector.

The "fit" column on Table 4 refers to the median explained variance, where the median is taken with respect to export destination countries. The simple model with exchange rates and a single common factor as regressors explains about 50% of the variation in the export price data. Most of the explained variation comes from the exchange rate effects. This can be read from the next to last column, where it is shown the median ratio of the variation due to exchange rate and the variation due to the common factor. The last column presents the fraction of countries residual series where the unit-root null is rejected. Critical values were obtained admitting three cointegrated variables. Given the low power of single-series tests, there is very strong evidence that residuals are stationary. This means the relationship summarized by the coefficient is not spurious.

Table 5. Panel cointegration results; one factor; only constant included *39 time periods*

			Alternative estimates			Final estimate (β)			
		Countries	ols	fm-ols	cup	fm-cup	fit	e/cmg	%Rej
High	Pharmaceuticals	29	-0.19	-0.13	-0.56	-0.56 (0.045)	56%	1.28	21%
	Medical, precision and optical instruments	32	-0.09	-0.05	-0.33	-0.32 (0.041)	53%	0.53	47%
Medium-high	Eletrical machinery and apparatus	39	-0.11	-0.27	-0.42	-0.42 (0.033)	52%	1.9	41%
	Machinery and equipment	51	-0.18	-0.3	-0.49	-0.5 (0.03)	55%	3.54	37%
	Chemicals excluding pharmaceuticals	43	-0.06	-0.51	-0.39	-0.4 (0.033)	74%	1.04	47%
	Motor vehicles, trailers and semi-trailers	41	-0.08	-0.44	-0.48	-0.48 (0.039)	71%	2.51	34%
Medium-low	Rubber and plastics products	53	-0.09	-0.24	-0.42	-0.38 (0.03)	55%	3.06	36%
	Non-metallic mineral products	53	-0.12	-0.36	-0.43	-0.38 (0.033)	63%	2.88	36%
	Basic metals and fabricated metal products	34	0.02	-0.54	-0.07	-0.07 (0.064)	78%	0.01	35%
Low	Food products, beverages and tobacco	46	-0.1	-0.36	-0.52	-0.46 (0.032)	74%	2.75	39%
	Wood, pulp, paper, paper products	48	-0.07	-0.61	-0.1	-0.09 (0.038)	88%	0.03	29%
	Other Manufacturing	33	-0.14	-0.21	-0.41	-0.36 (0.038)	60%	2.13	55%
	Textiles, leather and footwear	52	0.02	-0.37	-0.46	-0.47 (0.022)	71%	5.06	48%

Notes: (i) "ols" is the estimator from Kneeter. (ii) "fm-ols" is the estimator from Philips and Moon. (iii) "cup" is the continously updated estimator from Bai, Kai and Ng. (iv) "cup-fm" is the fully modified version of the last estimator, with standard error in parentheses. (v) The fit is the median explained variance among destination countries. (vi) The "e/cmg" is the median ratio of variances of the exchange rate effect and the marginal cost effect. (vii) "%Rej" is the fraction of rejections of the unit-root null aplied to model residuals of each country in the sector.

Table 5 shows results without the time-trend and can be similarly interpreted. Comparing with the previous table, the fully modified continuously updated estimator is larger in absolute value in most sectors but follows the previous estimates closely. There are only two abnormal sectors for which the full pass-through hypothesis cannot be rejected and for which the exchange rate effect contributes to very little of price variation. The Philips and Moon estimator does a much poorer job than before, indicating the greater importance of cross sectional dependency when no time-trend is included. Finally, the model residuals appear to be much less stationary then the previous case, suggesting that spurious regression may now be a serious issue. Overall, there is enough evidence to conclude that the time-trend specification is the superior one. As suggested before, it is likely that other permanent shock disturb the long-run relationship of interest, and the deterministic trend proxies for them. From this point on, we consider only the coefficients estimated under the assumption of country specific constant and time-trend.

Table 6. Sensitivity of parameter estimate (β) to the number of factors

			Number of Factors		
		Countries	1	2	3
High	Pharmaceuticals	29	-0.4	-0.57	-0.55
			(0.038)	(0.035)	(0.033)
	Medical, precision and optical instruments	32	-0.26	-0.3	-0.47
			(0.05)	(0.045)	(0.03)
Medium-high	Eletrical machinery and apparatus	39	-0.23	-0.25	-0.31
			(0.046)	(0.041)	(0.035)
	Machinery and equipment	51	-0.38	-0.42	-0.55
			(0.024)	(0.025)	(0.022)
	Chemicals excluding pharmaceuticals	43	-0.46	-0.44	-0.54
			(0.026)	(0.024)	(0.026)
	Motor vehicles, trailers and semi-trailers	41	-0.53	-0.47	-0.43
			(0.026)	(0.022)	(0.026)
Medium-low	Rubber and plastics products	53	-0.46	-0.49	-0.47
			(0.03)	(0.026)	(0.024)
	Non-metallic mineral products	53	-0.4	-0.48	-0.51
			(0.022)	(0.022)	(0.018)
	Basic metals and fabricated metal products	34	-0.53	-0.53	-0.65
			(0.039)	(0.031)	(0.037)
Low	Food products, beverages and tobacco	46	-0.46	-0.46	-0.47
			(0.024)	(0.025)	(0.02)
	Wood, pulp, paper, paper products	48	-0.59	-0.12	-0.08
			(0.024)	(0.032)	(0.032)
	Other Manufacturing	33	-0.38	-0.38	-0.3
			(0.033)	(0.027)	(0.025)
	Textiles, leather and footwear	52	-0.43	-0.49	0.08
			(0.02)	(0.02)	(0.034)

Note: Fully-modified estimates; constant and time trend included for all factor specifications

The inclusion of additional common factors preserves most of the results, apart from low technology sectors. Information criteria based on Bai and Ng (2002) provide no strong indication on the number of factors, with one or three factors being the preferred choice depending on the criteria; in any case, information criteria do not have robust properties in small finite samples. Table 6 reports the coefficient estimates and robust standard errors with as much as three common factors. Except for low technology ones where the single factor assumption seems to be essential, the pattern of results is fairly robust to the assumption. Without clear indication to the contrary by information criteria, parsimony and structural interpretation lead us to impose the single factor specification for all sectors[5].

[5] Additional robustness tests were not implemented due to the small sample size. In particular, parameter stability was not investigated, despite possible breaks around the shift from peg to float in the beginning of the sample. This issue should be addressed in future studies with more adequate samples.

5. Discussion

The results seem to improve on the previous panel literature, in the sense of a clearer picture on pricing-to-market behavior. Indeed, the long-run coefficients are much less dispersed then short-run analogues of the traditional panel literature reviewed before. Additionally, the sign pattern of negative mark-up effects significantly different from zero is more aligned with the aggregate evidence of incomplete pass-through which motivated the literature in the first place[6]. Moreover, in this section, we show the estimated coefficients have interesting patterns and connections with microeconomic fundamentals.

The industrial classification scheme used in the paper is essentially based on the research and development activities in each industry which may have connections with preference and market structure parameters. High technology sectors are constantly developing new product varieties to attend very specific consumer needs - for example, consumers of medical appliances and pharmaceuticals very often have few substitution possibilities. As for the market, competition is not expected to be high if measured by the average mark-ups in the industry which reflects large market shares among key participants. On the opposite extreme, low technology industries represent consolidated business with a large number of players offering very substitutable commodities. As a matter of fact, there is econometric evidence supporting this informal argument. For instance, Barroso (2009) found technological intensity is positively associated with product differentiation and supply elasticity.

Table 7 calculates simple averages of pass-through coefficients for each level of technological intensity. At least for this sample, it appears that pass-through is increasing with technology. It is suitable to add some theoretical underpinning to these observations. For that matter, is not hard to argue that high substitution and low shares lead to lower pass-through. In Dornbush (1987), high substitution is complementary to competitors' responses to expected aggregate price changes in the industry, and each firm is forced to adjust markups more strongly. Atkeson and Burstein (2008) show that mark-ups are sensitive to market-shares, the more so for higher elasticities of substitution. Since mark-shares reflect differences in marginal costs the result follows.

[6] Aggregate time-series studies for the Brazilian exports points to incomplete small pass-through. Ferreira and Sanso (1999) found long-run pass-through ranging from 10% to 27%; Tejada and Silva (2005) typical estimates ranges from 14% to 34%. Our disaggregated estimates are somewhat higher, possibly suggesting aggregation bias in time-series studies.

In both models, high market shares increases the influence on the sector price and therefore the pass-through that can be implemented. Of course, the actual details of the arguments leading to these interactions depend on the precise market structure and demand schedule assumed by the authors. But the pattern suggested by Table 7 and the underlying connection with microeconomic fundamentals argued for in the last paragraph both support models with this properties.

Table 7. Long-run margin and pass-through effects by technological intensity

		Mark-up	Pass-through		
High	Pharmaceuticals	-0,40	0,60		
	Medical, precision and optical instruments	-0,26	0,74	0,672	
Medium-high	Eletrical machinery and apparatus	-0,23	0,77		
	Machinery and equipment	-0,38	0,62		
	Chemicals excluding pharmaceuticals	-0,46	0,54		
	Motor vehicles, trailers and semi-trailers	-0,53	0,47	0,597	
Medium-low	Rubber and plastics products	-0,46	0,54		
	Non-metallic mineral products	-0,40	0,60		
	Basic metals and fabricated metal products	-0,53	0,47	0,539	
Low	Food products, beverages and tobacco	-0,46	0,54		
	Wood, pulp, paper, paper products	-0,59	0,41		
	Other Manufacturing	-0,38	0,62		
	Textiles, leather and footwear	-0,43	0,57	0,536	
					0,58

Notes: (i)The pass-through equals one plus the mark-up reduction. (ii) Summary measures are means of sector estimates by technological intensity.

Another interesting pattern is the negative relationship between the sector share in total manufacture exports and the degree of pass-through. Indeed, sector share explains 30% of the variance of estimated coefficients. This might be indicative of pooling bias that needs further investigation. Indeed, with fixed costs, domestic exporters would serve the most profitable destination markets first, where they can sustain a higher market share. But a sector with low participation in total exports is not highly developed and thus involves only the most profitable destinations. As a result, the pooled coefficient may give too much weight to low market share countries in traditional, low technology sectors. As a matter of fact, a similar argument could lead one to conclude that market share could be positive related to technological intensity. The possibility of pooling bias due to self-selection into destination countries should be further investigated. But care should be taken, because low market shares are also indicative of trade barriers which are of little consequence to the self-selection argument.

6. Conclusion

There is strong evidence of long-run pricing-to-market behavior by Brazilian exporters. Approximately 58% of an exchange rate appreciation would be passed-through to foreign consumer prices, with Brazilian exporters absorbing a 42% loss through reduced mark-ups. The degree of pass-through is positively related to the technological intensity of the industrial sector, a sensible pattern given the lower substitution between product varieties in high technology sectors. These results cannot be attributed to sticky-price constraints which by definition are not binding in the long-run. Therefore, the significant long-run effects support market structure explanations of incomplete pass-through and deviations from purchasing power parity, with possible normative consequences for trade and exchange rate policy.

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Appendix

This appendix shows how the cointegration equation can be derived from first principles, without, however, any pretense of generality. Indeed, as argued in the text, many other models would result in the same empirical specification.

Let residual demand in country i period t be $Q_{it}=Q_{it}(P_{it}E_{it})$ and exporter cost function be $C(\Sigma_i Q_{it})$, where prices p_{it} are set in the exporters currency and E_{it} is the bilateral exchange rate divided by the foreign price level. The objective is to maximize profits defined by $(\Sigma_i P_{it}Q_{it})$ - $C(\Sigma_i Q_{it})$ choosing prices. With the usual convexity and differentiability conditions, necessary and sufficient first order conditions define the pricing rule $P_{it} = \mu_{it}\, MgC_t$, where price is set equal to a markup over marginal cost. The markup $\mu_{it} = \varepsilon_{it} / (\varepsilon_{it} -1)$ is a function of demand price elasticities, which we assume not to be a constant. Therefore, the markup can be written as non constant a function of $\log(P_{it}E_{it})$. Marginal costs may be considered exogenous if the number of destinations is large, since, in this case, each destination demand will have a vanishing influence on marginal costs. This is why we dropped the destination index from the marginal cost above - in any case, estimation procedures in the text allowed for endogeneity correction using long run covariance of the innovations to the series. As another identifying assumption, let's suppose the markup time series resulting from this model is stationary, such that its long run average is $\mu_0>0$ corresponding to some point in its domain. Prices, marginal costs and exchange rates are otherwise unrestricted. Taking logs of the pricing rule and approximating the mark-up evaluated at its long run average, we get $p_{it} = d + \beta e_{it} + \lambda MgC_t$ where $\beta=\alpha/(1-\alpha)$ and $\lambda=1/(1-\alpha)$, with $\alpha=\mu_0'/\mu_0$.

This is essentially the empirical specification adopted in the paper. The difference is that equation (2) allows more flexibility in the form of (a) varying deterministic factors, (b) time and destination specific demand shocks and (c) destination specific proportionality factor to the common marginal costs to reflect destination specific features. This flexibility is meant to accommodate more general models and to provide a better fit to the data; but the structural interpretation is similar.

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