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Exchange Rate Misalignment and Its Effects on Agricultural Producer Support Estimates: Empirical Evidence from India and China

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ABSTRACT

There have been different degrees of exchange rate disequilibrium in the developing countries during transition or reform periods. The level of the exchange rate and its misalignment can have significant impacts on agricultural policy measures such as Producer Support Estimates (PSEs). In the conventional PSE analysis, however, the actual (nominal) exchange rates are used. There is general agreement that the use of actual exchange rates may introduce a bias in the PSE calculations, and that this bias can be substantial in some cases. But there is less agreement on the most appropriate alternative.

In this study, we utilize various time series techniques to derive estimates of the equilibrium exchange rates in India and China as determined by real economic fundamentals from the 1950s to the 2000s. The relevance and usefulness of the equilibrium exchange rates in the calculation of PSE for the two countries are then discussed. Drawing on the data sets and analyses developed earlier by Mullen, Orden and Gulati (2005) and Sun (2003), we find that agricultural support levels measured by the PSEs (from 1985-2002 for India and from 1995-2001 for China) are sensitive to alternative exchange rate assumptions. Specifically, exchange rate misalignments have either amplified or counteracted the direct effects from sectoral-specific policies. In India, such indirect effects are relatively small and mostly dominated by the direct effects. But in China, especially in recent years, the indirect effect from exchange rate misalignment (undervaluation) has been quite substantial.

Results from this study also show that the effect of the exchange rate depends on the relative importance of different PSE components. The increasing share of budgetary expenditures in India's total agricultural support in recent years has resulted in more pronounced exchange rate effects measured by commodity-specific percentage "PSEs" that use the value of production at international prices as the denominator compared to those measured by commodity-specific percentage Market Price Support (MPS) with the same denominator. For China, the exchange rate effects are more similar between the PSE and the MPS measures because budgetary expenditures have been relatively small.

The exchange rate effect when the PSE is "scaled up" from covered commodities to an estimate for the total agricultural sector is also demonstrated. Since the commodity coverage in both countries tends to be incomplete and the scaling-up procedure leads to a total MPS of greater magnitude, larger exchange rate effects are found in the scaled-up than the non-scaled-up version of total PSEs. The impact of scaling-up on the indirect effect is proportional to the share of covered commodities in the total value of agricultural production.

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1. Introduction

Agricultural policies in the developing countries play a very important role in determining domestic commodity prices and the returns to agriculture. The nature and degree of the policy interventions differ across countries thereby producing different types of impact on producers and consumers. Various agricultural policy indicators (APIs) such as the Producer Support Estimates (PSEs) have been constructed to evaluate and monitor these policy changes (Josling and Valdes, 2004). A problem with conventional analyses based on the APIs, however, is that they usually have a sector-specific focus that can miss the important linkages between economy-wide policies and the agricultural sector. By changing the relative prices of importables, exportables, and home goods, some economy-wide policies, such as the exchange rate policies, can have impacts on agricultural incentives that might overwhelm those from sectoral policies. The different effects of sectoral and economy-wide policies on agriculture in the developing countries were documented in a classic series of studies by Krueger, Schiff and Valdes (1988 and 1991).

The relevance of exchange rates in the PSE estimates has been pointed out by a number of authors including Harley (1996), Bojnec and Swinnen (1997), and Melyukhina (2002). This issue is particularly important for the developing countries since capital surges, macroeconomic instability and subsequent financial crisis in the last two decades, together with delayed or insufficient adjustments in the exchange rates, have generated substantial exchange rate misalignments in some of these countries. Pronounced misalignments in the exchange rate could potentially subsidize or tax the agricultural sector and lead to incorrect estimates of the level and sometimes the direction of agricultural support as measured by the PSE. In these cases, the effects of exchange rate misalignments have to be taken into account if meaningful calculations of the PSE are to be presented.

While there is general agreement that the use of misaligned exchange rates introduces a bias in the PSE calculations and that this bias can be substantial in some cases, there is much less agreement on the most appropriate alternative. The problem arises from the fact that it is fundamentally difficult to determine the equilibrium value of an exchange rate. A number of previous studies (e.g. Liefert, et al., 1996; Shick, 2002) have used effective exchange rates or purchasing power parity (PPP) exchange rates as the “equilibrium” exchange rates in their PSE calculations, and the findings from these studies generally indicate that alternative assumptions about exchange rates have significant impacts on the PSEs. Despite their plausible results, the calculations based on the PPP involve a high degree of discretion and the results are usually sensitive to the selection of a base year. Other models of equilibrium exchange rates are potentially preferred to the PPP approach in the PSE estimations (Harley, 1996).

More recently, the real equilibrium approach that relates the real exchange rate with economic fundamentals has gained prominence and been frequently used by both practitioners and policy makers to address issues of exchange rate misalignment and to test for over- or under-valued

currencies. The current study attempts to use the real equilibrium approach of exchange rate determination and identify the relevance and usefulness of the equilibrium exchange rate in the PSE calculations. We focus our analysis on two world's largest developing economies, India and China, where issues of exchange rate misalignment and the effects on the agricultural support have been important but nonetheless received little attention.

The paper is organized as follows. Section 2 empirically estimates the equilibrium exchange rates of India and China. Results in Section 2 are then applied to the MPS and PSE calculations in section 3. Summary and conclusions are provided in section 4.

2. Equilibrium Exchange Rate: India and China

To address the effects of exchange rate misalignments on agriculture support levels measured by the PSE, the first task is to establish the exchange rate equilibrium. There is no simple answer to what determined the equilibrium exchange rate, and estimating equilibrium exchange rates and the degree of exchange rate misalignment remains one of the most challenging empirical problems in open-economy macroeconomics (Williamson, 1994). The fundamental difficulty is that the equilibrium value of the exchange rate is not observable while the exchange rate misalignment refers to a situation in which a country's actual exchange rate deviates from such an unobservable equilibrium. The issue is further complicated by the fact that there exist a variety of models to determine the equilibrium exchange rates.¹

One strand of the empirical literature that originates from the real equilibrium model have gained prominence over the past decade and been frequently used in research and policy to address issues of exchange rate equilibrium and misalignment. Equilibrium exchange rates in this model are derived from a theoretical framework that maximizes representative agent's intertemporal utility subject to budget constraints (Agenor, 1998; Hinkle and Montiel, 1999). The reduced form solution of the model relates the long-run equilibrium real exchange rate to a set of steady-state values of supply-side, demand-side and policy variables (real economic fundamentals).² Since the early applications by Edwards (1989), a number of authors have used such approach to model exchange rates in the developing countries, e.g., Elbadawi, 1994; Elbadawi and Soto, 1994 and 1997; Mongardini, 1998; Baffes, Elbadawi and O'Connell, 1999; Rahman and Basher, 2001; Paiva, 2001; Zhang, 2001; Zhang, 2002; Cerra and Saxena, 2002; Mathisen, 2003 and Cady, 2003. These studies differ primarily by their selection of fundamentals and econometric methods.

This paper adopts the "reduced-form" real equilibrium approach where the real exchange rate is determined by the economic fundamentals. The set of fundamentals that may be identified (by theory) typically includes the following four categories: (1) Domestic supply-side factors and particularly the Balassa–Samuelson effect arising from faster productivity growth in the tradable good sector relative to the non-tradable good sector; (2) Fiscal policy, such as fiscal deficits as well as changes in the composition of government spending between tradable and non-tradable

¹ Commonly used approaches range from the simple Purchasing Power Parity (PPP) to more sophisticated models such as the Fundamental Equilibrium Exchange Rate (FEER) (Williamson, 1994), the Natural Real Exchange Rate (NATREX) (Stein, 1994), the Behavioral Equilibrium Exchange Rate (BEER) (Clark and MacDonald, 1999), etc.

² The BEER is similar to the specification of the real equilibrium model. But the theoretical underpinning for the BEER is the simple uncovered interest parity (UIP) condition, adjusted for the existence of a time varying risk premium (Clark and MacDonald, 1999).

goods; (3) International economic environment, including capital inflows, external transfers, and terms of trade; (4) Commercial policy such as trade liberalization in terms of a reduction in import tariffs and export subsidies. See Hinkle and Montiel (1999) for a discussion of the analytical model.

Given the nature of the underlying data series, the empirical analysis closely follows the contemporary non-stationary time-series modeling paradigm. First, the augmented Dickey-Fuller (ADF) tests are applied to the univariate data series to establish their temporal properties. Second, a Vector Error Correction Model (VECM) is specified and the Johansen cointegration method is used to determine if one or more long-run relationships exist among the system of variables. Third, structural hypothesis tests involving restrictions on cointegration coefficients as well as weak exogeneity are performed on the realized long-run relationships to determine uniquely the cointegrating vectors. Fourth, considering that the current values of the economic fundamentals themselves may not be at their long-run equilibrium level, the Hodrick-Prescott (H-P) filtering method is used to estimate long-run or permanent component of the fundamentals by removing the cyclical component from the data.³ Finally, the estimated cointegrating vector along with the permanent component of the economic fundamentals are used to calculate the equilibrium real exchange rates and gauge the corresponding misalignments.

Data Description

The sample period covers 1950-2003 for India and 1952-2002 for China, using annual data drawn from the *International Financial Statistics* of the IMF. A system of variables, $\mathbf{x}_t$, consisting of the real exchange rate and the underlying fundamentals, is formulated for each country with their constructions and composing data summarized in Table 1. Note that both the sample sizes and the choice of variables are dictated by data availability. All variables are in logarithmic forms except the interest rate.

The real exchange rate (*LRER*) is defined as the product of nominal exchange rate and the ratio of the consumer price indexes (*CPI*). While some other studies have focused on the *real effective exchange rate*, the real exchange rate defined here is a bilateral rate expressed in terms of domestic currencies per foreign currency (in this case the US dollar), so that an increase represents depreciation. The Balassa-Samuelson effect caused by differential productivity growth in the traded good vs. non-traded good sectors is approximated by the productivity change variable (*LPRO*). Consistent with the Balassa-Samuelson theory, an increase in the productivity in the tradable sector relative to the non-tradable sector appreciates the exchange rate, because it creates excess demand in the non-tradable sector. This variable is proxied by the log of annual growth of the industrial production index for India and the gross fixed capital formation for China (Cerra and Saxena, 2002; Zhang, 2001). The government expenditure variable (*LGEX*) as a percentage to the GDP is used to capture the effect of fiscal policies. Changes in the composition of government consumption affect the exchange rate in different ways, depending on whether the consumption is directed toward traded or non-traded goods. If an increase in government consumption is concentrated in non-traded goods, excess demand in this sector will

³ Alternatively, a moving average, a Beveridge-Nelson or a Gonzalo-Granger technique can be used to extract permanent component from the fundamentals (e.g. Elbadawi and Soto, 1997; Baffes, Elbadawi and O'Connell, 1999; Clark and MacDonald, 2000; Cerra and Saxena, 2002; and Mathisen, 2003).

lead to higher non-traded good price and thus real exchange rate appreciation. The opposite will happen if the government consumption is concentrated in traded goods.

Table 1: Variable Construction and Data Description

System of variables: $\mathbf{x} = [LRER, LPRO, LGEX, WIR, LTOT, LOPN]$	
Variable construction	
$LRER = \ln \left(e \cdot \frac{CPI^*}{CPI} \right)$	$WIR = NIR^* - \frac{CPI^* - CPI_{-1}^*}{CPI_{-1}^*}$
$LPRO = \ln \left(\frac{IPI}{IPI_{-1}} \right)$ (India)	$LTOT = \ln \left(\frac{XUV}{MUV} \right)$ (India)
$LPRO = \ln \left(\frac{FCF}{GDP} \right)$ (China)	$LTOT = \ln \left(\frac{VX}{VX_{-1}} \right)$ (China)
$LGEX = \ln \left(\frac{GEX}{GDP} \right)$	$LOPN = \ln \left(\frac{VX + VM}{GDP} \right)$
Description of data	
e	Nominal exchange rate
GEX	Government expenditure
NIR	Nominal interest rate
XUV	Export unit value (Index, 2000=100)
MUV	Import unit value (Index, 2000=100)
VX	Value of exports
VM	Value of imports
FCF	Fixed capital formation
CPI	Consumer price index (2000=100)
IPI	Industrial production index (2000=100)
GDP	Gross Domestic Product

Note: 1) * denotes foreign variables.

2) Variables $LPRO$ and $LTOT$ are constructed differently for India and China due to data constraints.

3) The data frequency is annual.

Source: IMF, IFS, various years.

To capture the change in international economic environment, the real world interest rate (WIR) is used, which is approximated by the US real interest rate calculated by subtracting the 1-year Treasury-Bill rate by the US inflation rate (percent change in the US CPI). It is widely accepted that the world interest rate interacts with capital flows into and out of the developing countries. A stylized effect associated with a reduction in the world interest rate and thus capital inflows is real exchange rate appreciations. However, Agenor (1998) and Hinkle and Montiel (1999) show

that the real exchange rate appreciation in the developing countries, following capital surges as a result of lower world interest rate, is only a temporary phenomenon, and in the long-run the real exchange rate tends to depreciate. The terms of trade (*LTOT*) is defined as the ratio of export price index to import price index for India. However, no consistent export and import price data are available for China and this variable is proxied instead by its export growth (one-lag difference of the log of exports) (Zhang, 2001). Early studies (e.g. Elbadawi and Soto, 1994) have shown that the effect of terms of trade is ambiguous. An improvement in the terms of trade increases national income which in turn increases demand for non-traded good leading to real exchange rate appreciation (an income effect). On the other hand, the improvement of terms of trade lowers the cost of imported inputs in the production of non-traded goods causing real exchange rate depreciation (a substitution effect). The openness (*LOPN*) is calculated as the ratio of the sum of imports plus exports to the GDP. Its use as a proxy for commercial policy is justified because of the difficulty of obtaining good time series data on import tariff and export subsidy and also because it may account not only for explicit commercial policy but also for implicit, though very important, factors such as quotas and exchange controls (Elbadawi and Soto, 1994).

The Unit Root and Cointegration Tests

Unit root and cointegration tests are conducted for the system of variables identified earlier. The augmented Dickey-Fuller (ADF) tests (Table 2) show that the null hypothesis of a unit root is accepted for all variables in levels except *LPRO* for India and *LTOT* for China, which seem to be trend stationary.⁴ The Phillips-Perron (PP) tests are also provided and they generally confirm the ADF test results. All series are then first differenced and the ADF and PP tests are re-conducted. For each series, both test statistics rose considerably, and the null hypothesis of a unit root is rejected at the 1 percent or 5 percent significance levels. It is thus concluded that all the variables in $\mathbf{x}_t$ are $I(1)$ in levels and $I(0)$ in differences.

If each series is an $I(1)$ process, common tests for the possibility of an equilibrium are the Johansen maximum likelihood method (Johansen, 1991). The Johansen test in this analysis starts with a specification of an unrestricted Vector Autoregressive model of order p ($VAR(p)$) which starts from a relatively small number of lags for each of the endogenous variables to economize the degrees of freedom:

$$\mathbf{x}_t = \mathbf{A}_1 \mathbf{x}_{t-1} + \mathbf{A}_2 \Delta \mathbf{x}_{t-2} + \dots + \mathbf{A}_p \mathbf{x}_{t-p} + \boldsymbol{\alpha} + \boldsymbol{\varepsilon}_t$$

where $\mathbf{x}_t$ is a $(n \times 1)$ vector of non-stationary $I(1)$ variables, $\mathbf{A}_i$ is a $(n \times n)$ coefficient matrix, $\boldsymbol{\alpha}$ is a $(n \times 1)$ intercept vector and $\boldsymbol{\varepsilon}_t$ is vector of error terms.⁵ The lag lengths of 2 and 1 are selected for India and China respectively based on a series of model diagnostic tests including the F-tests, the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC), and the Jarque-Bera and Ljung-Box tests for residuals (details are provided in Cheng, 2005).

⁴ The rejection of a unit root may result from small lag lengths, which adversely affect the size of the ADF and PP tests. When additional lags are included, the null hypothesis of a unit root cannot be rejected for these variables.

⁵ A vector of dummy variables is included in the VAR to take account of short-run shocks to the system. This vector contains three dummy variables for India representing two oil price shocks in the 1970s and the balance of payment crisis in the early 1990s, and three dummy variables for China representing the great famine during 1960-63 and two major exchange rate policy interventions in the early 1980s and mid 1990s.

Table 2: The Unit Root Test Results

Country Variable	Test Statistics					
	τ	$Z(\tau)$	τ_μ	$Z(\tau_\mu)$	τ_τ	$Z(\tau_\tau)$
India						
	Levels					
<i>LRER</i>	1.32	1.61	-0.30	-0.20	-2.15	-1.89
<i>LPRO</i>	-1.87	-1.75	-5.47**	-4.69**	-5.45*	-4.63**
<i>LGEX</i>	-2.09*	-2.40*	-2.13	-2.48	-2.27	-2.54
<i>WIR</i>	-0.70	-0.76	-1.64	-1.82	-1.07	-1.22
<i>LTOT</i>	-0.10	-0.55	-3.27*	-2.77	-3.55*	-2.83
<i>LOPN</i>	-0.60	-0.84	-0.26	-0.41	-2.38	-1.74
	First Differences					
<i>LRER</i>	-4.94**	-5.86**	-5.22**	-6.02**	-5.32**	-6.01**
<i>LPRO</i>	-4.27**	-24.74**	-4.21**	-24.48**	-4.15**	-25.28**
<i>LGEX</i>	-5.99**	-7.88**	-6.62**	-8.31**	-7.11**	-8.39**
<i>WIR</i>	-4.87**	-5.70**	-4.81**	-5.61**	-5.10**	-7.06**
<i>LTOT</i>	-4.91**	-9.73**	-4.86**	-9.55**	-4.87**	-13.51**
<i>LOPN</i>	-5.71**	-9.16**	-5.74**	-9.07**	-6.34**	-10.56**
China						
	Levels					
<i>LRER</i>	1.57	1.96	-0.29	-0.03	-2.32	-2.30
<i>LPRO</i>	-1.30	-2.30	-1.38	-2.34	-3.64*	-3.15
<i>LGEX</i>	-1.33	-1.96	-1.16	-0.91	-2.43	-2.50
<i>WIR</i>	-0.63	-0.74	-1.78	-1.78	-1.25	-1.17
<i>LTOT</i>	-2.98**	-2.93**	-4.42**	-3.80**	-4.49**	-3.73*
<i>LOPN</i>	-2.14*	-2.10*	0.20	0.92	-1.95	-2.06
	First Differences					
<i>LRER</i>	-4.67**	-5.40**	-5.28**	-5.81**	-5.27**	-5.77**
<i>LPRO</i>	-3.82**	-6.12**	-3.99**	-7.54**	-3.95*	-7.33**
<i>LGEX</i>	-2.41*	-7.10**	-3.38*	-7.64**	-3.74*	-7.50*
<i>WIR</i>	-4.56**	-5.25**	-4.50**	-5.16**	-4.75**	-6.15**
<i>LTOT</i>	-4.75**	-14.98**	-4.71**	14.61**	-4.67**	-14.27**
<i>LOPN</i>	-4.35**	-5.58**	-5.14**	-6.32**	-5.32**	-6.43**

Note: 1) **1% significance level, *5% significance level. Critical values are from MacKinnon (1991)

2) The test statistics τ , τ_μ , τ_τ and $Z(\tau)$, $Z(\tau_\mu)$, $Z(\tau_\tau)$ are ADF and PP tests respectively and correspond to three type of specifications: (i) no trend and no intercept; (ii) intercept only; and (iii) trend and intercept.

3) The lag length p of the ADF test is set by the AIC + 1 in every case. The PP test is based on Newey-West bandwidth selection using Bartlett kernel.

The Johansen test is conducted based on a VECM of the form:

$$\Delta \mathbf{x}_t = \Gamma_1 \Delta \mathbf{x}_{t-1} + \Gamma_2 \Delta \mathbf{x}_{t-2} + \dots + \Gamma_{p-1} \Delta \mathbf{x}_{t-p+1} + \Psi \mathbf{x}_{t-1} + \alpha + \varepsilon_t$$

where $\Gamma_i = -\sum_{j=i+1}^p \mathbf{A}_j$ and $\Psi = \sum_{i=1}^p \mathbf{A}_i - \mathbf{I}$ are $(n \times n)$ coefficient matrices. Alternatively, Ψ can be written as $\Psi = \mathbf{A}\mathbf{B}'$, for $\mathbf{A}$ an $(n \times h)$ matrix (adjustment coefficients) and $\mathbf{B}'$ an $(h \times n)$ matrix (cointegrating relationships). Table 3 shows the test results based sequentially on two model specifications with and without a deterministic trend in the data. At 5 percent significance level, the null hypothesis that $h = 0$ and $h \leq 1$ are rejected for India. However, the null hypothesis that the cointegrating rank is at most 2 is accepted. Hence, there is evidence that there are two cointegrating relationships among the variables ($h = 2$). For China, $h = 0$ is rejected and the cointegrating rank of at most 1 is accepted, so there is evidence of only one cointegrating relationship among the variables ($h = 1$).

Table 3: The Johansen Cointegration Test Results

Null Hypothesis	No deterministic trend in levels (I)				Deterministic trend in levels (II)			
	λ_{trace} test	λ_{trace} (0.95)	λ_{max} test	λ_{max} (0.95)	λ_{trace} test	λ_{trace} (0.95)	λ_{max} test	λ_{max} (0.95)
India	VEC(2) with dummies							
$h = 0$	151.55*	103.85	68.56*	40.96	147.86*	95.75	68.26*	40.08
$h \leq 1$	82.99*	76.97	38.82*	34.81	79.61*	69.82	38.32*	33.88
$h \leq 2$	44.18	54.08	20.60	28.59	41.29	47.86	20.60	27.58
$h \leq 3$	23.58	35.19	10.65	22.30	20.69	29.80	10.60	21.13
$h \leq 4$	12.92	20.26	9.60	15.89	10.09	15.49	9.23	14.26
$h \leq 5$	3.33	9.16	3.33	9.16	0.87	3.84	0.87	3.84
China	VEC(1) with dummies							
$h = 0$	119.04*	103.85	46.81*	40.96	108.43*	95.75	45.59*	40.08
$h \leq 1$	72.23	76.97	33.32	34.81	62.84	69.82	32.74	33.88
$h \leq 2$	38.91	54.08	14.26	28.59	30.10	47.86	13.77	27.58
$h \leq 3$	24.65	35.19	12.60	22.30	16.33	29.80	9.98	21.13
$h \leq 4$	12.05	20.26	9.20	15.89	6.35	15.49	5.92	14.26
$h \leq 5$	2.85	9.16	2.85	9.16	0.42	3.84	0.42	3.84

Note: 1) Case I: no intercept in the cointegrating equation or the VEC; Case II: intercept in the cointegrating equation and the VEC

2) h is the cointegrating rank.

3) * denotes rejection at 5% significance level.

The Structural Hypothesis Tests

A likelihood ratio (LR) test developed by Johansen and Juselius (1990) is used to test for restrictions on $\mathbf{B}$ and $\mathbf{A}$ individually and jointly. The restrictions in the long-run cointegrating relationships and the adjustment parameters are incorporated into the eigenvalue problem corresponding to each of the hypothesis to be tested. The LR test of the hypothesis is given by:

$$-2\ln\left(\hat{L}_R/\hat{L}_U\right) = -2\ln(Q) = T \sum_{i=1}^{\hat{h}} \ln \left\{ \frac{(1-\hat{\lambda}_i^*)}{(1-\hat{\lambda}_i)} \right\}$$

The test statistic is compared with a χ^2 distribution with the degrees of freedom equal to the number of parameter restrictions.

The test on $\mathbf{B}$ can be formulated as: $H_B : \mathbf{B} = (\mathbf{H}_1\varphi_1, \mathbf{H}_2\varphi_2, \dots, \mathbf{H}_h\varphi_h)$, where $\mathbf{H}_i$ is a $(n \times s_i)$ restriction matrix and φ_i is a $(s_i \times 1)$ vector of parameters to be estimated in the i th cointegrating relation ($i = 1, \dots, h$). In this formulation, there are s_i unrestricted parameters in the i th vector in $\mathbf{B}$ and $\mathbf{H}_i$ imposes k_i restrictions ($k_i + s_i = n$). While common restrictions placed on the cointegrating vectors include the exclusion and proportionality restrictions, we only consider the exclusion restrictions in this analysis. The restriction matrix $\mathbf{H}$ is constructed based on the following restrictions:

$$\mathbf{B}_{ij} = 0, \quad i = 1, \dots, 6, \quad j = 1, 2 \quad (\text{India})$$

$$\mathbf{B}_{i1} = 0, \quad i = 1, \dots, 6 \quad (\text{China})$$

The null hypotheses in each test and the LR test statistics are presented in Table 4. The test statistics are compared with a $\chi^2[r(n-s)]$ distribution (i.e., $\chi^2[2(6-5)]$ for India and $\chi^2[1(6-5)]$ for China).

Table 4: Exclusion Restrictions on B

Variable	India		China	
	Null Hypothesis	LR Statistic	Null Hypothesis	LR Statistic
<i>LRER</i>	$\mathbf{B}_{11} = \mathbf{B}_{12} = 0$	25.30*	$\mathbf{B}_{11} = 0$	4.21*
<i>LPRO</i>	$\mathbf{B}_{21} = \mathbf{B}_{22} = 0$	28.84*	$\mathbf{B}_{21} = 0$	0.04
<i>LGEX</i>	$\mathbf{B}_{31} = \mathbf{B}_{32} = 0$	23.24*	$\mathbf{B}_{31} = 0$	5.41*
<i>WIR</i>	$\mathbf{B}_{41} = \mathbf{B}_{42} = 0$	20.56*	$\mathbf{B}_{41} = 0$	5.49*
<i>LTOT</i>	$\mathbf{B}_{51} = \mathbf{B}_{52} = 0$	43.92*	$\mathbf{B}_{51} = 0$	7.41*
<i>LOPN</i>	$\mathbf{B}_{61} = \mathbf{B}_{62} = 0$	14.31*	$\mathbf{B}_{61} = 0$	4.09*

Note: 1) * denote rejection at the 5% significance level.

2) $\chi^2(2) = 5.99$ and $\chi^2(1) = 3.84$ at 5% significance level.

All the null hypotheses of zero coefficients in the cointegrating vectors ($\mathbf{B}$) are rejected for India, indicating that the real exchange rate and each of the fundamentals are relevant for the respective long-run relationships identified earlier. However, we fail to reject the null hypotheses that *LPRO* can be excluded from the long-run relationship for China.

The rejections of the null hypotheses relating to the long-run cointegrating coefficients in $\mathbf{B}$ for India warrant further attention. Since there are two cointegrating vectors for India and the above structural hypothesis tests on $\mathbf{B}$ are joint tests for coefficients in both vectors, rejection of the null hypothesis might indicate that each of the two cointegrating vectors represents a distinct

relationship in the cointegration space. To identify uniquely the equilibrium exchange rate relationship, we proceed by partitioning the system of variables into two “blocks” (the internal block: *LPRO* and *LGEX*; and the external block: *LRER*, *WIR*, *LTOT* and *LOPN*) and test the exclusion of both blocks in each cointegrating vector with the other unrestricted. The LR test statistic, distributed as χ^2 , indicates that at 5% significance level neither block can be excluded from the first cointegrating vector (with the second unrestricted). However, we fail to reject the exclusion of the “internal block” with $\chi^2(1) = 2.40$, and reject the exclusion of the “external block” with $\chi^2(3) = 21.71$ in the second cointegrating vector (with the first unrestricted). See Cheng (2005) for details.

To further investigate the model formation, a zero restriction on the adjustment parameters (**A**) or the weak exogeneity test is performed. The test on **A** can be formulated as:

$$H_A : \mathbf{A} = \mathbf{K}\tilde{\mathbf{A}}$$

where **K** is a $(n \times s)$ matrix, $\tilde{\mathbf{A}}$ is an $(s \times h)$ matrix of non-zero **A**-coefficients with $s \geq h$. The restriction matrix, **K**, is constructed based on the restrictions:

$$\mathbf{A}_{ij} = 0, \quad i = 1, \dots, 6, \quad j = 1, 2 \quad (\text{India})$$

$$\mathbf{A}_{i1} = 0, \quad i = 1, \dots, 6 \quad (\text{China})$$

The null hypotheses in each test and the test statistics are presented in Table 5. The test statistics are compared with a χ^2 distribution with the same degrees of freedom as the tests on **B**.

Table 5: Weak Exogeneity Tests on A

Variable	India		China	
	Null Hypothesis	LR Statistic	Null Hypothesis	LR Statistic
<i>LRER</i>	$\mathbf{A}_{11} = \mathbf{A}_{12} = 0$	25.48*	$\mathbf{A}_{11} = 0$	6.16*
<i>LPRO</i>	$\mathbf{A}_{21} = \mathbf{A}_{22} = 0$	16.92*	$\mathbf{A}_{21} = 0$	0.01
<i>LGEX</i>	$\mathbf{A}_{31} = \mathbf{A}_{32} = 0$	7.07*	$\mathbf{A}_{31} = 0$	3.69
<i>WIR</i>	$\mathbf{A}_{41} = \mathbf{A}_{42} = 0$	1.94	$\mathbf{A}_{41} = 0$	0.004
<i>LTOT</i>	$\mathbf{A}_{51} = \mathbf{A}_{52} = 0$	9.37*	$\mathbf{A}_{51} = 0$	1.31
<i>LOPN</i>	$\mathbf{A}_{61} = \mathbf{A}_{62} = 0$	13.57*	$\mathbf{A}_{61} = 0$	4.04*

Note: 1) * denotes rejection at the 5% significance level.

2) $\chi^2(2) = 5.99$ and $\chi^2(1) = 3.84$ at 5% significance level.

The null hypothesis of weak exogeneity of *WIR* cannot be rejected at the 5% significance level for India, while four variables (*LPRO*, *LGEX*, *WIR* and *LTOT*) are found to be weakly exogenous for China. A joint test of weak exogeneity for all four variables is also accepted with $\chi^2(4) = 6.47$ (p-value = 0.17). Thus *LPRO*, *LGEX*, *WIR*, and *LTOT* are concluded to be jointly weakly exogenous.

Combining all the information presented above, we pool the hypotheses H_B and H_A and test the restrictions on $\mathbf{B}$ and $\mathbf{A}$ jointly ($H_{B\&A} : H_B \cap H_A$). The restricted $\mathbf{B}$ and $\mathbf{A}$ for India and China take the following forms:

$$\text{India: } \mathbf{B}' = \begin{bmatrix} * & * & * & * & * & * \\ * & 0 & 0 & * & * & * \end{bmatrix} \text{ and } \mathbf{A}' = \begin{bmatrix} * & * & * & 0 & * & * \\ * & * & * & 0 & * & * \end{bmatrix}$$

$$\text{China: } \mathbf{B}' = \begin{bmatrix} * & 0 & * & * & * & * \end{bmatrix} \text{ and } \mathbf{A}' = \begin{bmatrix} * & 0 & 0 & 0 & 0 & * \end{bmatrix}$$

where the restrictions are imposed jointly on $\mathbf{B}$ as well as $\mathbf{A}$. Both joint restrictions cannot be rejected at 5% significance level with $\chi^2(3) = 4.44$ (p-value = 0.22) and $\chi^2(5) = 6.47$ (p-value = 0.26) for India and China respectively. On the basis of these results, Table 6 presents the restricted cointegration analysis for India and China.

Table 6: Restricted Cointegration Results

India						
B'						
<i>LRER</i>	<i>LPRO</i>	<i>LGEX</i>	<i>WIR</i>	<i>LTOT</i>	<i>LOPN</i>	Constant
1.000	12.450	-0.710	3.849	-0.505	-0.591	-4.557
	(1.325)	(0.077)	(1.088)	(0.182)	(0.075)	
-0.956	0.000	0.000	4.458	0.859	1.000	1.048
(0.242)			(1.661)	(0.397)		
A'						
0.009	0.015	-0.028	0.000	-0.020	0.006	--
(0.009)	(0.004)	(0.012)		(0.01)	(0.01)	
-0.010	0.000	0.000	0.000	0.015	-0.041	--
(0.009)				(0.011)	(0.014)	
China						
B'						
1.00	0.000	1.855	-11.735	1.913	-0.769	0.179
		(0.756)	(3.237)	(0.593)	(0.112)	
A'						
-0.031	0.000	0.000	0.000	0.000	-0.042	--
(0.009)					(0.015)	

Note: Standard errors are in parentheses.

Equilibrium Exchange Rates

Based on the restricted cointegration results in Table 6, the equilibrium real exchange rates (*LRER*) in India and China are calculated based on the following cointegrating equations with standard errors in parentheses:⁶

⁶ The calculations of the equilibrium exchange rates also draw on a recent study by Cheng and Orden (2005).

India:

$$LRER = 4.557 - 12.450LPRO + 0.710LGEX + 3.849WIR + 0.505LTOT + 0.591LOPN$$

(1.325) (0.077) (1.088) (0.182) (0.075)

China:

$$LRER = -0.179 - 0.000LPRO - 1.855LGEX + 11.735WIR - 1.913LTOT + 0.769LOPN$$

(0.000) (0.756) (3.237) (0.593) (0.112)

In general, the parameter estimates are consistent with expectations discussed earlier. The negative sign of the variable *LPRO* for India suggests that an increase in the productivity in the traded good sector relative to the non-traded good sector is associated with real exchange rate appreciation, which is consistent with the Balassa-Samuelson theory. The exclusion of *LPRO* in China's equilibrium exchange rate relationship indicates a lack of the Balassa-Samuelson effect in China. This is possibly due to the fact that China's domestic prices have been highly administered in the social economy setting, and the usual link between productivity and relative prices may be distorted leading to the break-down of the Balassa-Samuelson effect in this country. Similar findings are reported by Coudert and Couharde (2005), where the authors used a panel cointegration method based on a sample of 173 countries including China. An increase in the government expenditure (*LGEX*) causes the rupee to depreciate but the yuan to appreciate. The difference in the sign is due to the fact that government expenditures of India might have a higher content of traded good than that of China, which is consistent with Cerra and Saxena (2002) and Zhang (2001). The positive sign associated with *WIR* indicates that a reduction in the world interest rate depreciates the real exchange rates. This result is consistent with Agenor (1998) and Hinkle and Montiel (1999).

As shown earlier, the effect of the terms of trade on the equilibrium exchange rates can be ambiguous, depending on the relative importance between the substitution effect and the income effect. The positive signs on *LTOT* for India suggest the possible dominance of the substitution effect over the income effect and improvements in the terms of trade depreciate its currency. The reverse happens in the case of China where an increase in the terms of trade leads to real appreciation. The volume of trade, or degree of openness, as measured by the variable *LOPN* is an importance factor in determining the level of real exchange rate. In both countries, the positive signs confirm the findings in the literature that economic closedness is typically associated with overvaluation, and external liberalization aimed at reducing tariffs and eliminating trade restrictions causes currency depreciations.

In order to obtain the “long-run”, “steady state” or “permanent” values of the economic fundamentals in each country, the Hodrick-Prescott (HP) decomposing technique (Hodrick and Prescott, 1997) is applied. The HP filter decomposes the time series into a trend μ_t and stationary component $x_t - \mu_t$ by minimizing:

$$\sum_{t=1}^T (x_t - \mu_t)^2 + \lambda \sum_{t=2}^{T-1} [(\mu_{t+1} - \mu_t) - (\mu_t - \mu_{t-1})]^2,$$

where λ is an arbitrary constant reflecting the penalty of incorporating fluctuations into the trend. In this analysis, λ is chosen to be equal to 10 to match our annual data set (see Hodrick and Prescott, 1997). Based on the HP filtered values of the economic fundamentals as well as the

cointegrating vectors, the equilibrium exchange rate can be calculated for both countries. Figure 1 shows the equilibrium exchange rates as compared to the actual rates (reporting only 1985-2002).

Figure 1: The Actual and Equilibrium Real Exchange Rates

India

China

Source: Author's calculation

The real exchange rate of the Indian rupee is characterized by overvaluation as shown in Figure 1 (India). The overvaluation started in 1986 and ended in 1993, two years after the financial crisis in 1991. In the official descriptions of the events, India's exchange rate overvaluation and crisis has been attributed to continued current account deficits and reserve depletion, made worse by problems related to the Gulf War; and a loss of confidence in the government as political problems compounded the weak credibility associated with high fiscal deficits (Rangaranjan, 1993). Such explanation is consistent with the model results, by which weak economic fundamentals are associated with depreciated currency, and given insufficient devaluation of the actual exchange rate, overvaluation occurs.

The post-crisis adjustment program in India featured macroeconomic stabilization and structural reforms, especially in the direction of trade and financial liberalization. The effects of these measures were evident on the external sector. In the years after the crisis, rising capital inflows and shrinking trade deficits have led to continued accumulation of foreign exchange reserves. Large steps of devaluation in the early stage of the reform have caused the actual exchange rate to "overshoot" its long-run equilibrium in our model estimation, resulting in an undervaluation

for a number of years. In recent years, with a slight overvaluation, the rupee rate has remained close to its equilibrium values.

In comparison to the Indian rupee, the actual real exchange rate movements of the Chinese yuan during the post-reform era are in general characterized by undervaluations (Figure 1 (China)). It is important to note that, starting from 1999, there has been a period of sustained undervaluation. The widening gap between the equilibrium and the actual rate in recent years has stimulated a heated debate on the issue of undervaluation of the Chinese yuan. The US in particular has expressed serious concerns on the undervaluation in face of its exploding trading deficits with China, which amounted to over \$124 billion in 2003. A series of unfair trade practice petitions (“Section 301”) against China have been filed by a coalition of industry and labor groups. Despite the lobbying efforts from the US government along with strong supports from the IMF with the aim to eliminate China’s currency manipulation, the exchange rate of the Chinese currency remains fixed to the US dollar. The predicted degree of undervaluation in recent years (1999-2002) from this study averages about 20 percent which is broadly consistent with predictions from other recent empirical studies (see for example, Frankel, 2004; Goldstein, 2004; Coudert and Couharde, 2005).

3. Effects of Exchange Rates on the PSE

This section applies our model-based equilibrium exchange rate identified in the previous section to the PSE calculations.⁷ Building on the OECD’s PSE analytical framework, and drawing on recent PSE studies by Mullen, Orden and Gulati (2005) and Sun (2003), we examine the effects of exchange rate on the MPS and PSE for India and China.

Effects on MPS

The effects of exchange rate are directly reflected in the estimation of the MPS component of the PSE as an exchange rate is introduced to convert the world price into domestic currency. The MPS for a specific commodity (j) in monetary terms is defined as:

$$MPS_j = (P_j^d - P_j^{ar}) \times Q_j$$

where:

P_j^d : domestic producer price of commodity j

P_j^{ar} : adjusted reference price of commodity j

Q_j : quantity of commodity j

The adjusted reference price P_j^{ar} is the world market price expressed in domestic currency and adjusted by various costs. The cost adjustment differs depending on whether the commodity is an importable or an exportable, but in either case, the adjusted reference price can be expressed as:⁸

$$P_j^{ar} = P_j^{world} \times e + ADJ_j$$

⁷ The nominal equilibrium exchange rates used in the PSE calculations are derived from the corresponding real equilibrium rates by multiplying it with the ratio of the domestic CPI to US CPI.

⁸ Byerlee and Morris (1993) argue that the selection of a relevant reference price depends on the relationship between the autarky equilibrium price and the adjusted reference prices for imports and exports (see also Mullen, et al., 2004).

where:

P_j^{world} : world market price of commodity j

e : nominal exchange rate

ADJ_j : domestic cost adjustment for commodity j

The MPS_j in percentage terms is:

$$\%MPS_j = \left(\frac{P_j^d - P_j^{ar}}{P_j^{ar}} \right) \times 100$$

The $\%MPS_j$ is equivalent to the traditional nominal rate of protection (NRP).

Following the terminology of Krueger, Schiff and Valdes (1988), we define the “direct effect” that is induced by sector-specific policies as the $\%MPS_j$ calculated using the actual nominal exchange rate e :

$$\text{Direct Effect} = \%MPS_j(e) = \left[\frac{P_j^d - P_j^{ar}(e)}{P_j^{ar}(e)} \right] \times 100$$

where:

$P_j^{ar}(e)$: adjusted reference price of commodity j evaluated at e

The direct effect measures the proportionate excess of the domestic price from the adjusted reference price evaluated at the actual exchange rate. Similarly, the “total effect” that is induced by both sectoral and exchange rate policies is defined as the $\%MPS_j$ calculated using the equilibrium exchange rate e^* :⁹

$$\text{Total Effect} = \%MPS_j(e^*) = \left[\frac{P_j^d - P_j^{ar}(e^*)}{P_j^{ar}(e^*)} \right] \times 100$$

where:

e^* : nominal equilibrium exchange rate

$P_j^{ar}(e^*)$: adjusted reference price evaluated at e^*

The difference between the total and direct effect captures the “indirect effect” of misalignment of the exchange rate:

$$\text{Indirect Effect} = \%MPS_j(e^*) - \%MPS_j(e)$$

It is evident that the magnitude of the indirect effect is determined by the wedge between $\%MPS_j(e^*)$ and $\%MPS_j(e)$ caused by the exchange rate misalignment. When the exchange rate is overvalued, the indirect effect is negative; when the exchange rate is undervalued, the indirect effect is positive; and when there is no exchange rate misalignment, the indirect effect is zero (Cheng, 2005).

⁹ In all the calculations, we assume that there is no exchange rate pass-through so that the domestic prices are fixed. See Cheng (2005) for a relaxation of this assumption.

India

For India, the calculations are undertaken for 11 commodities including wheat, rice, corn, sorghum, sugar, groundnut, rapeseed, soybeans, and sunflower, chickpeas, and cotton. Following Mullen, Orden and Gulati (2005), the $\%MPS_j$ for six commodities (wheat, rice corn, sorghum, groundnuts and sugar) are calculated using the Byerlee and Morris (1993) procedure in determining the reference prices (see footnote 7), while rapeseed, soybeans, sunflower, chickpeas and cotton are assumed to be importable for all years. Table 7 shows the direct, indirect and total effects measured by $\%MPS_j$ for India for the period 1985-2002.

The sample period 1985-2002 is divided into four distinct subperiods for the presentation of our results. Period one (I) represents the sustained overvaluation period from 1985 to 1989. Period two (II) is the crisis period from 1990 to 1992 when the exchange rate was overvalued but under active adjustment. Period three (III) covers the slight undervaluation period of 1993-1997. The last period (IV) is the stable exchange rate period from 1998-2002 when the actual exchange rate is close to the equilibrium rate with a slight overvaluation. Average exchange rate misalignment for each period is also presented in Table 7.

The numbers on the direct effect are equivalent to the conventional measures of $\%MPS$ or NRP. On average, the agricultural protection or disprotection measured by the direct effect has shown a counter-cyclical pattern in India. Specifically, the direct effect was generally positive when world commodity prices were low during the first period (I). It became almost neutral during the crisis years (period II), and turned to disprotection when the world prices strengthened in the mid-1990s (period III). The world prices have since then followed a downward trend, and in the most recent period (IV) an unweighted average indicates a slight disprotection. These results are consistent with Mullen, Orden and Gulati (2005), who report their commodity $\%MPS_j$ on an annual basis.

The indirect effect caused by exchange rate misalignments has had quite different impacts on India's agriculture in comparison to the direct effect. On average, India's agricultural sector has been indirectly penalized by exchange rate overvaluation in periods I, II and IV, but subsidized by exchange rate undervaluation in period III. The averages in Table 7 also show that the effect of the exchange rate counteracted the direct effect in periods I, II and III, and reinforced the direct effect, on average, in period IV. The indirect effect was greater in the years before and during the crisis, when the exchange rate was continuously misaligned. In the post-crisis years, as the result of decreased magnitude of exchange rate misalignment following macroeconomic restructuring, the indirect effect has dampened down. Noticeably, the indirect effect of the exchange rate is smaller in absolute value than the direct effect in periods I and III, indicating the dominance of sectoral-specific policies over economy-wide policies (exchange rate in this case). The opposite happened in periods II and IV. This result is somewhat different from that of Krueger, Schiff and Valdes (1988 and 1991) in which they found that the economy-wide policies such as the exchange rate play a more dominant role across a range of developing countries (not including India and China) in an earlier period up to the mid-1980s.

Table 7: Direct, Indirect and Total Effects by %MPS_j, India

Commodity	1985-1989 (I)			1990-1992 (II)			1993-1997 (III)			1998-2002 (IV)		
	% Misalignment -8.7 ^a			% Misalignment -6.9			% Misalignment 4.3			% Misalignment -3.7		
	Direct	Indirect	Total	Direct	Indirect	Total	Direct	Indirect	Total	Direct	Indirect	Total
Wheat	-0.7	-4.1	-4.8	-18.3	-4.7	-23.0	-20.2	4.4	-15.8	11.7	-1.1	10.6
Rice	-10.2	-5.8	-16.0	-30.7	-5.0	-35.7	-32.4	3.2	-29.2	-2.3	-3.3	-5.6
Corn	20.4	-7.9	12.5	-0.7	-3.5	-4.2	-5.4	1.7	-3.7	-0.6	-0.8	-1.5
Sorghum	41.5	-4.7	36.8	13.7	-3.3	10.4	9.2	0.9	10.1	15.0	-4.0	10.9
Groundnuts	65.2	-11.8	53.4	14.0	-1.5	12.4	2.7	2.0	4.7	9.2	-3.4	5.8
Sugar	61.8	-12.8	49.0	-3.5	-4.1	-7.6	-4.1	4.1	0.0	24.7	-4.6	20.1
Rapeseed	18.9	-9.0	9.9	23.8	-6.6	17.2	-4.3	3.9	-0.3	-4.8	-3.6	-8.4
Soybeans	-1.7	-7.9	-9.7	-8.2	-5.4	-13.6	-20.6	3.1	-17.5	-30.3	-2.7	-33.0
Sunflower	37.5	-11.3	26.1	29.9	-7.8	22.2	-10.3	3.6	-6.7	-19.0	-3.2	-22.2
Chickpeas	3.5	-8.9	-5.4	-8.9	-6.2	-15.1	-31.2	2.9	-28.4	-3.5	-4.2	-7.7
Cotton	19.8	-13.0	6.8	-1.2	-7.3	-8.5	-11.0	4.5	-6.5	-13.0	-4.1	-17.1
Average ^b	23.3	-8.8	14.4	0.9	-5.0	-4.1	-11.6	3.1	-8.5	-1.2	-3.2	-4.4

Note: a. Average of annual percentage misalignment.

b. Simple unweighted average.

Source: Authors' calculation.

China

For China, we base our evaluation on the PSE analysis by Sun (2003), which is limited to the years 1995-2001. The % MPS_j is based on 9 commodities: wheat, rice, corn, sorghum, peanut, cotton, rapeseed, soybeans and sugar. Due to data constraints, we do not apply the Byerlee and Morris (1993) procedure in China: rice, corn, sorghum and peanut are assumed to be exportable and wheat, cotton, soybeans, rapeseed and sugar are assumed to be importable. For presentation of the results, the sample period is divided in the two subperiods (I and II), representing, respectively, periods of slight currency overvaluation (1995-1998) and more intense undervaluation (1999-2001). Table 8 shows the direct, indirect and total effect by commodity-specific % MPS_j in China.

Table 8: Direct, Indirect and Total Effect by % MPS_j , China

Commodity	1995-1998 (I)			1999-2001 (II)		
	%Misalignment -5.8 ^a			%Misalignment 15.8		
	Direct	Indirect	Total	Direct	Indirect	Total
Rice	-10.0	-0.9	-10.9	1.7	11.4	13.1
Corn	7.1	-1.0	6.1	5.5	11.7	17.2
Sorghum	-18.1	-0.9	-19.0	-32.6	7.6	-25.0
Peanut	-39.1	-0.8	-39.9	-39.2	6.7	-32.5
Wheat	3.8	-1.4	2.4	-18.2	9.0	-9.2
Cotton	-7.6	-1.2	-8.8	-28.0	8.1	-19.9
Soybeans	18.9	-1.5	17.4	15.5	12.7	28.2
Rapeseed	11.5	-1.5	10.0	2.8	11.2	14.0
Sugar	31.0	-1.7	29.3	15.1	12.8	27.9
Average ^b	-0.3	-1.2	-1.5	-8.6	10.1	1.5

Note: a. Average of annual percentage misalignment.

b. Simple unweighted average.

Source: Authors' calculation.

Except for rice, the direct effect of price interventions measured by % MPS_j has shown decreased protection or increase disprotection on the representative commodities in 1999-2001 compared to 1995-1998. The commodity-specific results are similar to those reported in Sun (2003) and Mullen, et al. (2004).¹⁰ The unweighted average disprotection rate increased from -0.3 in period I to -8.6 in period II.¹¹ As demonstrated for India, exchange rate misalignment in China has either indirectly taxed or subsidized the agricultural sector. The indirect effect of exchange rate overvaluation is relatively small and averages about -1.2 percent in period I. However, the exchange rate undervaluation in period II has had a much greater impact on the agricultural sector. It indirectly subsidized agricultural prices by 10.3 percent, on average, counteracting the direct effect and resulting in a positive total effect of 1.5 percent.

¹⁰ Mullen, et al. (2004) report the results only for five commodities (wheat, soybeans, sugar, rice and corn).

¹¹ In a longer term context, there appears to be a move toward lessened disprotection of agriculture in China (see Mullen et al. (2004) and Cheng and Sun (1998).

Effects on Commodity-Specific PSE

The calculation of commodity-specific PSE requires that budgetary payments be allocated across commodities to determine the budgetary support for a given product, BP_j , where “j” denotes a specific commodity. Once budgetary payments are allocated among commodities, the product-specific PSE_j is the sum of the MPS_j and BP_j . Similar to the MPS_j , the PSE_j measures can be expressed on a percentage basis. Following Mullen, et al. (2004), we define a measure of $\%PSE_j$ (using a “trade economist’s denominator”) that expresses support received by farmers as a percentage of the value of their output at adjusted reference price:

$$\%PSE_j^{ar} = \left(\frac{MPS_j + BP_j}{VOP_j^{ar}} \right) \times 100$$

where:

$$VOP_j^{ar}: \text{value of production of commodity } j \text{ at adjusted reference price} = P_j^{ar} \times Q_j$$

Corresponding, the direct, total and indirect effects are given as:

$$\text{Direct Effect} = \%PSE_j^{ar}(e) = \left[\frac{MPS_j(e) + BP_j}{VOP_j^{ar}} \right] \times 100$$

$$\text{Total Effect} = \%PSE_j^{ar}(e^*) = \left[\frac{MPS_j(e^*) + BP_j}{VOP_j^{ar}} \right] \times 100$$

$$\text{Indirect Effect} = \%PSE_j^{ar}(e^*) - \%PSE_j^{ar}(e)$$

It can be seen from the above definitions that the direct, total and indirect effects measured by $\%PSE_j^{ar}$ are different from those measured by $\%MPS_j$. When the budgetary payment is positive, the direct and total effects of $\%PSE_j^{ar}$ indicate more protection (or less disprotection). However, the impact of inclusion of budgetary payment on the indirect effect depends on the direction of exchange rate misalignment. As for the indirect $\%MPS$ effect itself, if the exchange rate is overvalued, the indirect effect is negative (i.e. shows less protection or more disprotection). If the exchange rate is undervalued, the indirect effects is positive (shows more protection or less disprotection). See Cheng (2005) for details.

India

Using the same commodity set and sample periods as in the $\%MPS_j$, Table 9 presents the direct, indirect and total effects by commodity specific $\%PSE_j^{ar}$ for India. On average, the direct effect on representative commodities exhibits the same counter-cyclical pattern as found in $\%MPS_j$: support to agriculture drops from a relatively high level in period I to near neutrality in period III, and then rose in period IV. However, the direct effect measured by the $\%PSE_j^{ar}$ indicates more agricultural protection than the $\%MPS_j$, due to the inclusion of positive budgetary payments in the calculation. As the budgetary payments in India become more important in the aggregate support to farmers, their impacts on the measured support levels increase. For instance, the difference between the unweighted average direct effect measured by $\%PSE_j^{ar}$ and $\%MPS_j$ during period IV is more than 17 percent.

Table 9: Direct, Indirect and Total Effect by $\%PSE_j^{ar}$, India

Commodity	1985-1989 (I)			1990-1992 (II)			1993-1997 (III)			1998-2002 (IV)		
	%Misalignment -8.7 ^a			%Misalignment -6.9			%Misalignment 4.3			%Misalignment -3.7		
	Direct	Indirect	Total	Direct	Indirect	Total	Direct	Indirect	Total	Direct	Indirect	Total
Wheat	15.6	-4.7	10.9	5.6	-6.1	-0.5	3.4	5.7	9.1	45.0	-1.5	43.5
Rice	-1.7	-6.5	-8.1	-20.1	-5.7	-25.9	-21.0	3.7	-17.3	15.9	-4.2	13.7
Corn	27.5	-8.5	19.1	7.7	-3.9	3.8	4.2	1.9	6.1	11.6	-0.9	10.6
Sorghum	45.7	-4.8	40.9	20.7	-3.5	17.2	18.3	0.9	19.2	27.9	-4.3	23.5
Groundnuts	69.8	-12.2	57.7	18.4	-1.6	16.8	8.8	2.2	11.0	20.0	-3.7	16.4
Sugar	71.8	-13.6	58.2	4.5	-4.4	0.1	5.3	4.5	9.8	37.4	-5.1	32.4
Rapeseed	30.0	-9.9	20.1	35.9	-7.3	28.5	8.0	4.4	12.4	16.8	-4.4	12.4
Soybeans	6.4	-8.6	-2.2	-3.2	-5.7	-8.8	-17.2	3.2	-14.0	-25.2	-2.9	-28.1
Sunflower	46.5	-12.2	35.3	35.5	-8.1	27.4	-4.0	3.8	-0.2	-6.1	-3.6	-9.7
Chickpeas	8.2	-9.3	-1.1	-1.1	-6.7	-7.8	-23.6	3.2	-20.4	8.4	-4.7	3.6
Cotton	58.0	-17.9	40.1	30.6	-10.0	20.6	16.7	7.3	24.0	27.1	-6.2	20.9
Average ^b	34.3	-9.8	24.6	12.2	-5.7	6.5	- 0.1	3.7	3.6	16.3	-3.8	12.6

Note: a. Average of annual percentage misalignment.

b. Simple unweighted average.

Source: Author's calculation

The inclusion of budgetary payment in $\%PSE_j^{ar}$ has also made the effect of exchange rate misalignment larger for each commodity in each period than for the $\%MPS_j$. The indirect effect shows more disprotection during periods of overvaluation (I, II and IV), and more protection during period of undervaluation (III). In each period, the indirect effect of exchange rate misalignment again counteracts the direct effect. Specifically, exchange rate overvaluation has indirectly taxed agriculture in periods I, II and IV, but undervaluation has subsidized agriculture in period III. The indirect effect is, on average, less than the direct effect in absolute values, reflecting again the more dominant role of sectoral policies (except period III).

China

Table 10 shows the direct, indirect and total effects measured by commodity-specific $\%PSE_j^{ar}$ in China. It is important to note that there have been little explicit policy instruments for direct payments or subsidies to farmers in China (Sun, 2003). Instead, there have been various taxes and fees by the local and central governments targeted at specific agricultural commodities. Following Sun (2003), these taxes and fees were treated as negative payments to farmers.¹² Nonetheless, the budgetary payments including the taxes and fees represent only a small proportion of aggregate support to the farmers. When allocated to each representative commodity according to its share in the total value of production, they are less than 10 percent of the price support (MPS) in most cases (Sun, 2003).

Table 10: Direct, Indirect and Total Effect by $\%PSE_j^{ar}$, China

Commodity	1995-1998 (I)			1999-2001 (II)		
	%Misalignment -5.8 ^a			%Misalignment 15.8		
	Direct	Indirect	Total	Direct	Indirect	Total
Rice	-9.9	-0.9	-10.8	2.5	11.4	13.9
Corn	7.2	-1.1	6.1	6.8	11.7	18.5
Sorghum	-18.2	-0.8	-19.0	-31.7	7.5	-24.2
Peanut	-39.0	-0.8	-39.8	-38.5	6.8	-31.7
Wheat	3.8	-1.3	2.5	-17.2	9.0	-8.2
Cotton	-7.5	-1.2	-8.7	-27.2	8.2	-19.0
Soybeans	19.1	-1.6	17.5	16.9	12.7	29.6
Rapeseed	11.6	-1.4	10.2	4.0	11.3	15.3
Sugar	31.0	-1.7	29.3	15.2	12.8	28.0
Average ^b	-0.2	-1.2	-1.4	-7.7	10.2	2.5

Note: a. Average of annual percentage misalignment.

b. Simple unweighted average.

Source: Authors' calculation.

Because of the dominance of the MPS component in the aggregate support, the inclusion of small budgetary payments in the calculation of $\%PSE_j^{ar}$ has had little impact on the direct, total

¹² Starting in 2000, China launched its pilot reforms on rural tax and fee system, and one of the reforms is to lower, exempt or abolish the taxes on the farmers (Sun, 2003). However, these reform measures were still in their experimental stages during 2000-2001 and their impacts were little on this analysis.

and indirect effects in China. Similar to the $\%MPS_j$, the direct effect of agricultural policies on unweighted average has disprotected the farmers of these covered commodities for the whole period and the rate of disprotection increased from -0.2 in 1995-1998 to -7.7 in 1999-2002. The indirect effect of exchange rate misalignment has amplified the direct effect in the first period but counteracted the direct effect in the second period. The indirect effect is more pronounced in the second period due to greater exchange rate undervaluation, which indirectly subsidized the agricultural sector by 10.2 percent. Furthermore, the small budgetary payments (or taxes) have also made the change of indirect effect over time primarily determined by the degree of China's exchange rate misalignment.

Effects on Total PSE

The total PSE expressed in nominal terms for all agricultural producers is the sum of total MPS and aggregate budgetary payments. The calculation of total MPS, according to the OECD approach consists of three steps. First, a nominal value of MPS is estimated for individual products, the set of which is known as the covered “MPS commodities.” The second step is to sum the product-specific MPS results into an MPS_c for the covered commodities. In the third step the MPS_c for covered commodities is “scaled up” to all products based on the share (k) of the covered commodities in the total value of agricultural production. The final step or “MPS extrapolation procedure” can be expressed as $MPS = MPS_c/k$, where MPS is the estimated total market price support. With the scaling-up, the OECD “Total PSE” is calculated as $PSE = MPS + BP$. Without the scaling-up the total PSE is $PSE_c = MPS_c + BP$. Similar to commodity-specific PSEs, total PSE measures can be expressed on a percentage basis (denoted by $\%PSE$) using $(VOP + BP)$ or VOP^{dr} as the denominator, where VOP and VOP^{dr} are the total value of agricultural production at domestic producer prices and world reference prices, respectively, and BP is the total budgetary payments. In the following analysis, the total $\%PSE$ is calculated using the OECD denominator $(VOP+BP)$ as commonly reported.

Again, we define the direct, total and indirect effect in terms of $\%PSE$. The three effects in the non-scaled-up case can be adapted from the $\%PSE_j$ by replacing the MPS_j and $(VOP_j + BP_j)$ with MPS_c and $(VOP + BP)$. With scaling-up, the effects are given as:

$$\text{Direct Effect} = \%PSE(e) = \left[\frac{MPS_c(e)/k + BP}{VOP + BP} \right] \times 100$$

$$\text{Total Effect} = \%PSE(e^*) = \left[\frac{MPS_c(e^*)/k + BP}{VOP + BP} \right] \times 100$$

$$\text{Indirect Effect} = \%PSE(e^*) - \%PSE(e) = \frac{1}{k} \left[\frac{MPS_c(e^*) - MPS_c(e)}{VOP + BP} \right] \times 100$$

It is evident that the direct, indirect and total effects in the scaled-up version are different from those in the non-scaled-up version. The impact of scaling-up on the support level (direct and total effects) depends on the sign of MPS: if MPS is negative, then the scaling-up will indicate more disprotection and if MPS is positive, the scaling-up will indicate more protection. The scaling-up can lead to different degrees of change in the direct and total effects depending on the relative magnitude of MPS and BP. However, the scaling-up magnifies the indirect effect by exactly

$1/k$, the inverse of the share of covered commodities in the total value of agricultural production.

India

Table 11 shows the three effects in India without and with scaling-up. The covered commodities are the same as those used in the previous calculations. In general, the support to agriculture (direct and total effect), scaled-up or non-scaled-up, indicates a similar counter-cyclical pattern, rising when world prices are low (as in periods I and IV) and falling when world prices strengthen (as in period II and III).¹³

Table 11: Direct, Indirect and Total Effect by %PSE, India

	Non-scaling-up (%PSE _c)			Scaling-up (%PSE)		
	Direct	Indirect	Total	Direct	Indirect	Total
1985-1989 (I)	6.8	-3.0	3.8	9.7	-6.4	3.3
1990-1992 (II)	-0.9	-3.3	-4.1	-9.4	-7.0	-16.4
1993-1997 (III)	-4.4	2.1	-2.3	-19.5	5.0	-14.5
1998-2002 (IV)	8.2	-1.1	7.1	8.8	-2.5	6.3

Source: Authors' calculation.

In period I and IV, the MPS_c component of the PSE is relatively small with the same sign as the budgetary payment, and the scaling-up does not have a significant impact on support levels indicated by direct and total effects. In contrast, when the MPS_c is more important and of similar magnitude but opposite signs as the budgetary payment in period II and III, the scaled-up direct and total effects are several times larger than the non-scaled-up ones. The direct effect decreased more than 10 times from -0.9 to -9.4 in period II. The change in the direct effect is smaller in period III, dropping from -4.4 to -19.5. The scaling-up also causes the total effect to decrease from -4.1 to -16.4 in period II, and from -2.3 to -14.5 in period III. The impact of scaling-up on the direct and total effects in India also depends on the sign of MPS. In terms of the direct effect, where the MPS is evaluated at the actual exchange rate, the scaling-up has led to higher protection rates in periods I and IV when the MPS is positive, but higher disprotection rates in periods II and III when the MPS is negative. The results are different for the total effect since the MPS in this case is evaluated at the equilibrium exchange rate.

The scaled-up and non-scaled-up indirect effects show different impact of exchange rate misalignment on India's agriculture during different periods. The exchange rate effects are consistent with the previous commodity-specific analysis: when the exchange rate was overvalued (periods I, II and IV), the indirect effect works against the agricultural sector. The opposite happens when the exchange rate was undervalued in period III. The scaling-up has had a uniform impact on the indirect effect for each period, which more than doubles in the scaled-up than the non-scaled-up case. The reason for this is that the share of covered commodities in total value of production is about 0.45 for each of the periods.

¹³ Again, the results are very similar to those reported by Mullen, Orden and Gulati (2005).

China

Table 12 shows the direct, indirect and total effect in China without and with the scaling-up. The scaled-up or non-scaled-up direct effect measured by total PSE indicates that China's support to the agriculture has remained negative in each period and level of discrimination against agriculture increased over the short time period covered. This result is somewhat different from Sun (2003) and Mullen, et al. (2004) who report the annual %PSE based on different sets of commodities. The indirect effect of exchange rate shows slight disprotection in period I but greater protection in period II.

Table 12: Direct, Indirect and Total Effect by %PSE, China

	Non-scaling-up (%PSE _c)			Scaling-up (%PSE)		
	Direct	Indirect	Total	Direct	Indirect	Total
1995-1997 (I)	-0.4	-0.9	-1.3	-0.8	-1.9	-2.7
1998-2001 (II)	-0.7	7.0	6.3	-3.0	16.8	13.8

Source: Authors' calculation.

In contrast to India, the MPS_c (evaluated at the actual exchange rate) in China is a more important component of PSE than the budgetary payments for most of the years. In period I when the MPS_c is significantly larger than the budgetary payment, the impact of scaling-up on the direct effect is small (twice as large in the scaled-up relative to the non-scaled-up version). In period II, the decrease in the magnitude of MPS_c raises the impacts of scaling-up on the direct effect to more than 4 times.

Since the MPS_c evaluated at the actual exchange rate is negative on average in each period, the scaling-up has uniformly led to more disprotection indicated by the direct effect. However, when the MPS_c becomes positive when calculated using the equilibrium exchange rate in period II, the total effect with scaling-up indicates more protection than that without scaling-up. Similar to India, the scaling-up in China has had a uniform impact on the indirect effect, which is about twice as large as the non-scaled-up counterpart. This again corresponds to the fact that the share of covered commodities in total value of production averages about 0.5 in the two periods.

4. Summary and Conclusions

The level of the exchange rate and its disequilibrium can have significant impacts on the agricultural sector. Although there are repeated claims and indeed widespread agreement that exchange rate misalignments can lead to inaccurate calculations of the support measures, empirical studies on the issue have been scant and mainly focused on the effects of exchanges rate on the nominal rate of protection (NRP). There have been attempts to consider the role the exchange rate plays in more comprehensive agricultural policy indicators such as the PSEs, but the calculations usually use simple adjustment approaches based on effective or PPP exchange rates. In this analysis, we propose an alternative approach to determining equilibrium exchange rates and apply the measures in evaluating the MPS and the PSE.

Nominal equilibrium exchange rates corresponding to estimated real equilibrium rates are applied to the MPS and the PSE calculations in the two countries based on earlier analyses by Mullen, Orden and Gulati (2005) and Sun (2003). Our results indicate that the indirect effect of

exchange rate overvaluation has potentially taxed the agricultural sector in India during the period of 1985-1992 and 1998-2002. However, the magnitude of these indirect effects is smaller in the later periods when the actual exchange rate moves closer to its equilibrium value. The indirect effect of exchange rate misalignment is, in general, smaller than the direct effect from sectoral-specific policies. For China, the indirect effect of exchange rate undervaluation in the period of 1999-2001 has had a much greater impact on the measured support level than overvaluation in 1995-1998. It has more than offset the disprotection to the agriculture and led to positive a total effect.

The relative importance of different PSE components also affects the estimates of the direct, indirect and total effects. First, the inclusion of budgetary payments in the analysis changes the effects of agricultural and exchange rate policies measured by the MPS. Commodity-specific PSEs show that such changes are more pronounced in India where the budgetary payments have become a more important component of the PSE. The direct and total effects by the commodity-specific PSEs indicate more protection and less disprotection. But the impact of including budgetary payment on the indirect effect can be different depending on the direction of exchange rate misalignment: more disprotection when the exchange rate is overvalued and more protection when the exchange rate is undervalued. In contrast to India, China has had little explicit budgetary payments to farmers and the effects measured by the PSE do not differ much from those measured by the MPS.

Second, the scaling-up procedure has had important impacts on the different effects measured by the total PSE. In India, the scaling-up leads to more significant changes in the direct and total effects during periods when the MPS becomes larger in magnitude than the budgetary payments. In contrast, changes in the direct and total effects in China are similar across periods since the MPS plays a more dominant role in the total PSE in most of the years. The effect of scaling-up on support levels depends on the sign of MPS: if MPS is negative, then scaling-up leads to more disprotection and if MPS is positive, the scaling-up leads to more protection. In addition, the impact of scaling-up on the indirect effect is directly related to the share of covered commodities in the total value of agricultural production. In India and China, the scaling-up leads to greater indirect effects which double the non-scaled-up counterparts.

In general, the estimation results are consistent with previous studies. However, the results in this study are highly dependent on the modeling strategy of equilibrium exchange rates. Alternative models of equilibrium exchange rates exist and it would be ideal to conduct a sensitivity analysis of the PSE where results from different approaches to equilibrium exchange rates are compared. In addition, the above measures of different effects assume that the domestic price of the relevant commodity remains unchanged even though the exchange rates have been changed. In other words, this study, like many others assessing exchange rate effects on the NPR, ignores the “pass-through” of exchange rates on domestic prices. Recent research efforts have been directed toward addressing these considerations (see Cheng, 2005).

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