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Income Inequality and the Real Exchange Rate: Linkages and Evidence*

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A simple model with non-homothetic preferences and purchasing-power parity for tradables shows that improved income inequality decreases the price of nontradables, resulting in a real depreciation. This hypothesized negative association between income inequality and the real exchange rate has robust empirical support from random- and fixed-effects models, dynamic panel estimations, and panel vector autoregressions. Thus, policies that improve a country's income distribution, by leading to a depreciation of the real exchange rate, may improve the competitiveness of its goods. However, this income inequality-real exchange rate relationship does not imply that dramatic redistributive policies will automatically bring about a real depreciation of the domestic currency.

Key Words: Income distribution; Gini coefficient; Price of nontradables; Real exchange rate; Real depreciation.
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1. INTRODUCTION

Many attempts have been made to understand why poverty reduction remains so "elusive" despite the best-intentioned policies and increased globalization (Porto, 2008, p. 179). It is important to understand the ways in which globalization impacts poverty, including the specific transmission mechanisms. Trade is clearly a critical linkage between countries and the real exchange rate is critically important to trade. While the real exchange rate's impact on sustainable export growth has been widely studied, little work has been done on whether and how a country's income distribution affects its real exchange rate. By developing a simple model in which preferences are assumed to be non-homothetic and purchasing-power parity holds for tradables, we find that an improved (or more equal) income distribution decreases the price of nontradables, resulting in a real depreciation of the exchange rate. Our theoretical finding of a clear negative relationship between income inequality and the real exchange rate contrasts with the theoretical model of Garcia (1999) in which there can be either a positive or negative relationship between these two variables. It is important to understand this effect because if an improved income distribution is negatively related to the real exchange rate then through its positive impact on the trade balance it may accelerate economic growth, ceteris paribus. Establishing the linkage between income inequality and real exchange rates helps us understand one of the ways in which a more equal income distribution might be beneficial to growth (Aghion et al., 1999, p. 1619).

In our empirical work, unlike previous researchers, we control for business-cycle effects and liquidity. We provide evidence in support of a robust negative association between income inequality and real exchange rates. This negative relationship implies that policy makers should be concerned with the distributional implications of their policies — not only for social and political reasons, but also because income inequality has long-run effects on the real exchange rate through changes in the price of nontradables. Policies that improve a country's income distribution, by leading to a depreciation of the real exchange rate, may also improve the competitiveness of its goods and thus its trade-balance performance.

After a brief discussion of the most relevant previous work, we present our simple model that links income inequality to the real exchange rate. This model shows that changes in income inequality affect the price of nontradables, which in turn affect the real exchange rate. Empirical evidence provided by fixed- and random-effects models as well as panel vector-autoregressions supports these conceptual arguments. The long-run negative association between income inequality and the real exchange rate is large, significant, and robust to alternative specifications of the reduced-form model and estimation methodologies. The concluding section offers some policy recommendations.

2. PREVIOUS WORK

Although much work has been done on trade policy and economic growth[1] and on income inequality and economic growth or development, relatively little attention has been paid to the relationship between income inequality and trade policy. Here we focus on one macroeconomic aspect of this relatively neglected but important relationship by examining how a country's level of income inequality affects its real exchange rate and thus its trade balance, ceteris paribus. By postulating one possible transmission mechanism from income inequality to the real exchange rate, our work also provides a nice complement to that of Agenor (2004), who showed that real exchange-rate depreciation increases the welfare of the poor. It also helps to better understand why there is a negative association between growth and inequality or as Easterly (2007, p. 773) puts it, "higher inequality hinders development."

As a nice example of the importance of trade policy to economic growth, Rodrik (2008) develops a model for the linkages between the real exchange rate and the rate of economic growth. According to Rodrik (2008), currency undervaluations (a high real exchange rate) are found to stimulate economic growth, especially for developing countries; the operative channel appears to be the size of the tradable sector. He also finds that the relative price of tradable goods to nontradable goods[2] — that is, the real exchange rate — seems to play a more fundamental role in the convergence of developing country with developed country incomes. So there seems to be an important connection between the real exchange rate and development at work, although no explicit connection is made by Rodrik between the real exchange rate and income inequality.

Although Kuznets (1955) was the first to make the connection between inequality and development, Easterly (2007) is a nice example of more recent work that has been done on the relationship between income inequality and economic growth. Using cross-country data he finds that agricultural endowments predict inequality and inequality predicts development, con-

[1]See, e.g., Rodriguez and Rodrik, 2000 and Rodrik, 2008.

[2]Studies on the relative price of tradables to nontradables can be grouped into three interrelated categories. The first focuses on the relationship between national price levels and the relative price of tradables and nontradables (see e.g., Kravis and Lipsey, 1988; Bergstrand, 1991; and De Gregorio et al., 1994). The second focuses on the relative price of tradables to nontradables as a possible source of errors in purchasing power parity (see e.g., Kim, 1990; Davutyan and Pippenger, 1985; and Rogoff, 1992). The third focuses on the co-movement of real exchange rates and the relative price of tradables to nontradables (see e.g., Strauss, 1999 and Kakkar and Ogaki, 1999).

cluding that high inequality is a "large and statistically significant barrier to prosperity." Clarke (1995) finds that income inequality is negatively, and robustly, correlated with long-run growth. Although both these papers do an excellent job of deciphering the connections between income inequality and growth, any connection between trade issues and policies (including the real exchange rate) and income inequality is not discussed.

Mitra and Trindade (2005) take a microeconomic approach to examine the role of income inequality in the determination of trade flows and patterns. They use a simple Heckscher-Ohlin framework and find that countries with identical factor endowments and technology gain from trade if their degrees of inequality are different. Dalgin et al. (2008) extend this microeconomic approach and find that income inequality is an important determinant of import demand: while imports of luxury goods increase with the importing country's inequality, imports of necessity goods decrease with it.

However, to the best of our knowledge there are only two papers that highlight macroeconomic factors in the linkages between income inequality and the real exchange rate. Kocherlakota and Pistaferri (2007) demonstrate empirically that after controlling for several possible additional determinants there is a statistically significant relationship between real exchange rate growth and cross-country differences in inequality growth. However, because they use both consumption- and income-distribution data, they actually examine the relationship between inequality (technically, the "consumption distribution") and real exchange rates rather than between the income distribution and real exchange rates, which is what we do in this paper. In addition, their critical finding is that a worsening of the distribution (or an increase in the right-tail income share, $LQ5$ in our study) decreases the price of nontradables. However, many researchers find the opposite result — that a worsening of the income distribution should increase the price of nontradables.[3] Finally, they find that right-tail but not left-tail inequality growth rates are significant in affecting real exchange-rate growth. But our study shows that both the left-tail ($LQ1$) and right-tail ($LQ5$) income shares are important determinants of the real exchange rate.[4]

While Garcia (1999) investigates the relationship between the income distribution and real exchange rates using a Salter-Swan type general equilibrium model with heterogeneous agents (non-homothetic preferences), his

[3]For example, De Gregorio et al. (1994, p. 1229) claim that tradables has a less-than-unitary income elasticity of demand, whereas that for nontradables exceeds unity. Moreover, Bergstrand (1991, p. 333) shows that, under nonhomothetic tastes, higher real per capita income will raise the price of nontradables. Anderson (1987, p. 201) and Lluch et al. (1977) also claim that demand for nontradables tends to be income elastic.

[4]Their analysis also consists solely of regressions (with and without year and region dummies), whereas our analysis consists of both fixed- and random-effects models as well as vector-autoregressions.

findings about this relationship are ambiguous as it can be either positive or negative. To investigate this relationship we develop a simple model that makes three assumptions: non-homothetic preferences; purchasing-power parity holds for tradables; and the classical definition of the real exchange rate — that it is the relative price of tradables to nontradables. Our model yields an unambiguous testable hypothesis: that reduced income inequality decreases the price of nontradables, resulting in a real depreciation. Moreover, our empirical model has broader control variables than Garcia's (e.g., liquid liabilities and fixed capital formation, a proxy for business-cycle effects), both of which turn out to be highly significant. Furthermore, our period of analysis, 1980-2007, is more recent than Garcia's, 1965-1990. In addition, our estimation methodologies include fixed- and random-effect models, panel vector autoregression analysis, and dynamic panel estimations (system-GMM). Our estimation results are generally robust under different specifications of the real exchange rate equation and using different estimation models.

Before moving on to our empirical work, we present the linkages between the income distribution and the real exchange rate: an improved (more equal) income distribution (a lower Gini coefficient) implies a decreased demand for nontradables so the real exchange rate will increase (a real depreciation).

3. LINKAGES BETWEEN THE INCOME DISTRIBUTION AND REAL EXCHANGE RATE

Consider an economy that produces two composite goods, tradables and nontradables. This economy is composed of two heterogeneous income groups, a high income group and a low income group, that have the same income share.

ASSUMPTION 1. *Non-homothetic preferences between different income groups.*

Following other researchers, e.g., Fajgelbaum et al. (2011) and Mitra and Trindade (2005, p. 1254), we assume that the total demand for goods depends not only on aggregate income, but also on the distribution of that income.[5] As Dalgin et al. (2008, p. 747) state: "We begin our argument with the empirical fact that tastes cannot be considered to be homothetic." Empirical evidence supporting this assumption can be found in Dalgin et al. (2008), Haq and Meilke (2007), and Tchamourliyski (2002).

[5]See Mitra and Trindade (2005, p. 1254) for more details about the assumption of non-homothetic preferences and its history in the trade literature. On p. 1255 they provide a list of papers that provide empirical support for this assumption.

If we also assume that prices of nontradables are flexible, which is a useful benchmark if current-account adjustment occurs gradually, then the income elasticity of demand for nontradables is higher for rich households than for poor households. Many studies make this assumption or provide empirical support for it; a recent example is Larrain (2010, p. 796), who claims that higher income increases the demand for nontradables.[6]

Although evidence in favor of purchasing-power parity (PPP) is sensitive to the choice of the base country, time period, and type of tests used (see, e.g., Kim, 1990; Davutyan and Pippenger, 1985; Rogoff, 1992; Oh, 1996; Wu, 1996; and Lothian, 1997), we follow Taylor (2009, p. 3) who in his summary of 18 empirical studies finds "strong and robust support for long-run PPP." Thus, we make the following assumption:

ASSUMPTION 2. *Purchasing power parity holds only for tradables,*

$$e + P_T^* - P_T = 0 \quad so \quad e = P_T - P_T^* \tag{1}$$

where e is the nominal exchange rate, which represents the home currency price of the foreign currency, P_T^* and P_T are the foreign and domestic price of tradables, respectively, and * denotes the foreign economy. All variables are expressed in logs.

PROPOSITION 1. *If income inequality decreases (increases), ceteris paribus, then the real exchange rate depreciates (appreciates).*

Equation (2) presents the real exchange rate, REX, as an implicit function:

$$REX = f(\overline{P_T}, P_{NT}(G)) \tag{2}$$

[6]Giles and Hampton (1985) found that the estimated higher-income groups' total expenditure elasticities for housing, household operation, and transportation were consistently higher than those of lower-income groups in New Zealand. Using data from Australia, the United Kingdom, and the United States, they also found that elasticities of demand for transportation were higher in the higher income countries. Bergstrand (1991) shows that increased income, under non-homothetic preferences, can lead to a shift in the demand from tradables to nontradables, resulting in an increase in the price of nontradables. Using data from the United States on rents, house values, and housing characteristics, Hansen et al. (1996, p. 175) find evidence of substantial variation in income elasticities for housing demand across income classes; the income elasticity for housing demand increased monotonically with income for both renters and owners. Chinn (2000) and Samuelson (1964), using the Penn effect, claimed that higher levels of income are associated with a greater demand for nontradables (such as services) and thus with a higher relative price of nontradables.

where P_{NT} is the price of nontradables, which is a function of the income distribution, G, the Gini coefficient.[7] $\overline{P_T}$ is the price of tradables, which will be determined in the world tradables market. Following Magee and Magee (2008) we assume this price is the same for all countries, so it is independent of a country's income distribution.[8] To examine the effect of the Gini coefficient on the real exchange rate, we need to look at the components of the right-hand side of Equation (3):

$$\frac{dREX}{dG} = \frac{\partial REX}{\partial P_{NT}} \frac{dP_{NT}}{dG} < 0 \tag{3}$$

Assumption 1 means $\frac{dP_{NT}}{dG} > 0$ because an improved (more equal) income distribution — a lower Gini — implies a decreased demand for nontradables since the high-income group has a higher elasticity of demand for nontradables than does the low-income group. Ceteris paribus, there will be a decrease in P_{NT}. It is interesting to note that this result is consistent with findings from the literature on transition economies. Svejnar (2002, p. 19) presents data on Gini coefficients for central and eastern European countries that shows inequality increasing (Ginis rising) following the collapse of the Soviet system in the late 1980s. Egert (2002, pp. 9-10) notes that nontradable prices are significantly lower in less developed countries, so with the "catch-up process," these countries will experience an increase in nontraded goods prices. If this catch-up process occurs as it did in the former Soviet bloc countries then rising income inequality (an increasing Gini) would be associated with increases in P_{NT}. De Gregorio et al. (1994) examine the inflation experience of 14 OECD countries over the period 1970-1985; they find that the relative price of nontradables increased for these countries over this time period, when their real GDP per capita were still "converging" (p. 1237).

We assume that home and foreign prices are weighted averages of the prices of tradables and nontradables, with weights $(1 - \varphi)$ and φ, respec-

[7] We use the Gini coefficient as a measure of income inequality. The income shares of the lowest and highest quartiles are used as supplements to the Gini coefficient.

[8] Although it is often assumed that large countries' actions can affect the price of tradables, we are assuming that all countries face an exogenous price of tradables. This assumption is supported by the empirical work of Magee and Magee (2008) that shows that even the United States is a small country in world trade in that its trade policies have a negligible effect on world prices. Kravis and Lipsey (1988, p. 475) also find support for their hypothesis that prices of tradables though higher in richer countries are much more similar among countries than prices of nontradables.

tively:[9]

$$P = (1-\varphi)P_T + \varphi P_{NT} \tag{4}$$
$$P^* = (1-\varphi^*)P_T^* + \varphi^* P_{NT}^* \tag{5}$$

Defining the real exchange rate, $REX = e + P^* - P$, and using Equations (1), (4) and (5) we have

$$REX = (e + P_T^* - P_T) - \varphi(P_{NT} - P_T) + \varphi^*(P_{NT}^* - P_T^*), \tag{6}$$

which, if we recall that $e + P_T^* - P_T = 0$ from Equation (1), gives us:

$$REX = -\varphi(P_{NT} - P_T) + \varphi^*(P_{NT}^* - P_T^*) \tag{7}$$

Partially differentiating Equation (7) with respect to P_{NT} yields

$$\frac{\partial REX}{\partial P_{NT}} < 0$$

Looking again at Equation (3) we can now see that an improved (more equal) income distribution (a lower Gini) will be associated with an increase in its REX, a real depreciation. Before testing this hypothesis using several models (fixed-effects, random-effects, system GMM, and panel vector autoregressions), we first present the estimating equations and data and discuss the correlations between key variables.

4. ESTIMATING EQUATIONS AND DATA

Because our focus is on the relationship between the real exchange rate (REX) and income inequality ($Income\ distribution$), the fundamental regression takes the form:

$$REX_{i,t} = a_{i,t} + b[Income\ distribution]_{i,t} + c[Control\ variables]_{i,t} + e_{i,t} \tag{8}$$

where i represents a specific country and t represents a particular year. In order to reduce endogeneity problems, we use (log) real exchange rate, REX, as the dependent variable rather than some measure of inequality as there is no generally accepted theoretical or empirical evidence as to what determines cross-country differences in inequality.[10]

[9] This assumption is quite common (see, e.g., Chinn, 2006, p. 117) and, because the share of the manufacturing sector in GDP does not vary much for most countries over long time horizons, is consistent with the data. Strauss (1999), e.g., showed that for most economies over a 30-year time period the share of the manufacturing sector in GDP varied from only 1-3.5%.

[10] Tanzi (1998) provides some descriptive arguments on what determines income inequality but his hypotheses do not have any theoretical or empirical support.

Following many other researchers we define the dependent variable, the real exchange rate REX, as the relative price of tradable to nontradable goods (see, e.g., Frenkel and Mussa, 1985; Zietz, 1996; and Edwards, 1989).[11] Note that this means the REX[12] is not an asset price; rather it reflects the relative attractiveness and thus competitiveness of a country's goods, so our *Control variables* will focus on other variables deemed critical to a country's international competitiveness in goods. As Zietz's (1996, p. 158) empirical work shows, the relative price of nontradables is a "key long-run driving force" behind the (U.S.) trade balance.

Income distribution is represented either by $GINI$ coefficients or the lowest quintile, $LQ1$, or highest quintile, $LQ5$, of income share. We use panel estimation methodology as it can handle more comprehensive problems induced by country-specific effects and any potential endogeneity problems.

Admittedly the literature on the determinants of real exchange rates is vast; we include as our macroeconomic *Control variables* only those that have been most often identified in the literature as having a long-run stable relationship with the real exchange rate and are critical to the competitiveness of a country's goods (see, e.g., Faruqee, 1995; Chung and Kang, 2005; and Chung et al., 2009). These variables are: the terms of trade, TOT; liquid liability to GDP, LLY; real GDP per capita, $RCGDP$; a country's openness, $OPEN$; manufacturing sector productivity, $FXKY$, which is a proxy for business cycles; and a country's stock of human capital, SCH.

Equation (9) presents the reduced-form model for the real exchange rate that includes as explanatory variables a measure of income equality (e.g., the Gini coefficient, $GINI$) as well as our *Control variables*:

$$\begin{aligned} REX \;=\; & \beta_0 + \beta_1 \log(GINI) + \beta_2 \log(TOT) + \beta_3 \log(LLY) \\ & + \beta_4 \log(RCGDP) + \beta_5 \log(FXKY) + \beta_6 \log(OPEN) \\ & + \beta_7 \log(SCH) + \varepsilon \end{aligned} \tag{9}$$

[11]The REX is defined as pi/pc where pi, the price of tradables, is the price level of investment, and pc, the price of nontradables is the price level of consumption. Both price series are from the Penn World Table 6.3 so they are internationally comparable; in addition, these price series are the best proxies available for the price of tradables and price of nontradables in the Penn World Table. Feenstra (1996) shows that (at least for the United States over the period 1972-94), most imports were for Investment rather than Consumption. Also, the Penn World Table price data for different countries indicates that "price level of investment" is the best proxy for the price of tradables. The pc data are superior to the CPI as the CPI series contains many traded goods. Chinn (2006, pp. 116-122) has a nice discussion of the various approaches used to define the real exchange rate and the pros and cons of different empirical measures for REX.

[12]We define REX as the relative price of tradables to nontradables since we are interested in how the change in the price of nontradables affects the real exchange rate. In other words, we use Equation (2) as the empirical definition of the real exchange rate.

The relationship between the terms of trade, TOT, and the real exchange rate is the result of income and substitution effects that depend on the source of the terms-of-trade variation. The likely result is that a deterioration in the terms of trade leads to a real depreciation, whereas an improvement in the terms of trade will lead to an appreciation (see, e.g., Edwards and van Wijnbergen, 1987; Neary, 1988; Kahn and Ostry, 1991; and Tokarick, 1995). Because the TOT is the relative price of the exportable and REX is the relative price of tradables, we need an estimation technique that treats all the variables in the model as endogenous; this is done in our panel vector autoregressions below.

Liquid liability adjusted by GDP, LLY, is a measure of liquidity and serves as a proxy for the inflationary pressures in the economy. The likely effect of LLY on REX is positive because increased liquidity in an economy will cause a real depreciation.

$RCGDP$ is real GDP per capita. Although Kravis and Lipsey (1988, p. 476) find that an increase in $RCGDP$ is associated with a decrease in the relative price of tradables so REX would fall, a real appreciation, it is also reasonable to expect that an increase in $RCGDP$ might fall more heavily on tradables than nontradables for lower income countries, driving up the relative price of tradables so that REX rises.

$OPEN$, a country's degree of trade openness to international trade, is captured by exports plus imports divided by GDP; it serves as a proxy for increased integration and thus heightened global competition. The likely effect of $OPEN$ on REX is negative because, ceteris paribus, increased amounts of exports and imports should be associated with a lower relative price of tradables, a lower REX.

$FXKY$ is public investment on fixed capital divided by GDP. $FXKY$ is used as a proxy for the business cycle and is expected to have a negative effect on REX as typically during expansions productivity improves and this is associated with a real appreciation (see, e.g., De Gregorio et al., 1994, p. 1238).

SCH is human capital measured by years of schooling. The likely effect of SCH in our REX regression is positive since increased education is associated with a lower price of nontradables (Alcala and Ciccone, 2004).[13]

To estimate regional and income level effects, we use several regional dummy variables: $EASIA$ for East Asia, $LATIN$ for Latin America, and $OECD$ for the OECD member countries.

Our data are discussed briefly below; more details are provided in the Appendix. The key variables in our analysis are the Gini coefficients, which are taken from the UNU-WIDER World Income Inequality Database ($WIID$)

[13]Garcia (1999) uses human capital as a control variable.

V2.0c (2008). From the $WIID$ database we chose 69 countries over the period 1980-2007. While this database provides Gini coefficients for more countries and a longer time period, we chose these 69 countries as the quality of the estimates for them was acceptable for our analysis.[14] The period 1980-2007 was selected as it was the only time period that provided comparable data for our panel regressions. Easterly (2007, p. 761) has a nice discussion about the flaws of international inequality datasets; however, he chooses an earlier version of the $WIID$ dataset citing the procedure used to remove bias due to survey methodology. Table 1 presents our most important data series, Gini coefficients and real exchange rates, for the 69 countries in our sample; the values presented are 28-year averages.

As supplements to the Gini coefficients, we also use the lowest ($LQ1$) and highest ($LQ5$) quintiles of income shares, which are reported as 28-year averages in Table 2. These data are also from the UNU-WIDER $WIID$ (2008).

Table 3 presents descriptive statistics for all our variables except $LQ1$ and $LQ5$.[15] Other than the inequality data, all data are from the Penn World Table 6.3, IMF's International Financial Statistics CD-Rom, and the World Development Report (2010). Panel A shows statistics for the full sample; Panel B is for the OECD countries; and Panel C is for the non-OECD countries. It is interesting to note that the mean value of the Gini coefficient for the OECD countries is much smaller than that of the non-OECD countries.

5. CORRELATION ANALYSES

Before we move on to the regressions, correlation analyses between the Gini coefficients, the price of nontradables, and the real exchange rates are performed to test the validity of the various assumptions used in the model. Since the main contribution of this paper is identifying a new transmission mechanism from the income distribution to the real exchange rate through a change in the price of nontradables, we perform simple correlation analyses

[14]Given the various sources of the Gini coefficients, we used the following criteria: availability, quality (provided by the source), and coverage (area, population, and age).

[15]Panel unit root tests for the full sample indicate that all data except $GINI$ are stationary (Im et al., 2003) and additional panel cointegration tests show that all series are cointegrated.

Variables	$\log(OPEN)$	$\log(FKXY)$	$\log(RCGDP)$	$\log(LLY)$	$\log(TOT)$	$\log(GINI)$	$\log(REX)$	$\log(SCH)$
$Z_{\%tbar}$	-23.03^{**}	-20.60^{**}	-23.96^{**}	-20.31^{**}	-20.17^{**}	0.06	-22.66^{**}	-2.37^{**}
Cointegration test (Pedroni, 2004) for $\log(REX)$, $\log(GINI)$, $\log(OPEN)$, $\log(LLY)$, $\log(TOT)$ and $\log(SCH)$								
Group rho test	3.536^{**}	Group Philips-Perron test		-2.691^{**}	Group ADF test		-2.466^{*}	

TABLE 1.

Countries' mean real exchange rates and Gini coefficients, 1980-2007

Country	Real Exchange Rate REX	Gini Coefficients $GINI$	Country	Real Exchange Rate REX	Gini Coefficients $GINI$
Algeria	0.61	37.65	Lesotho	1.86	57.63
Australia	0.83	33.59	Malawi	2.72	50.1
Austria	0.82	25.03	Malaysia	1.18	45.17
Bangladesh	1.88	35.5	Mauritania	1.68	45.18
Belgium	0.74	27.34	Mauritius	2.41	37.9
Bolivia	1.58	56.62	Mexico	0.96	52.55
Brazil	1.26	58.64	Morocco	2.11	39.17
Cameroon	3.26	47.92	Nepal	1.05	38.52
Canada	0.81	28.88	Netherlands	0.9	31.75
Chile	0.93	54.59	New Zealand	0.92	35.63
China	1.42	35.59	Nigeria	2.25	52.13
Colombia	1.58	57.43	Norway	0.66	34
Costa Rica	0.78	48.34	Pakistan	1.43	37.25
Cote d'Ivoire	2.98	39.85	Panama	0.76	55.89
Denmark	0.76	38.67	Peru	1.1	52.1
Dom. Rep.	0.98	50.47	Philippines	1.33	47.44
Fiji	1.55	49	Poland	0.92	29.66
Finland	0.69	27.64	Portugal	0.77	37.18
France	0.8	28.52	Rwanda	4.62	37.16
Germany	0.84	30.55	Singapore	0.68	46.24
Ghana	2.6	35.4	South Africa	2.08	47
Greece	0.78	34.08	Spain	0.74	32.32
Guatemala	0.82	52.7	Sri Lanka	1.59	45.28
Honduras	0.85	53.99	Sweden	0.9	25.39
Hong Kong	1.03	43.4	Switzerland	0.75	30.63
Hungary	1.19	24.4	Taiwan	1.14	30.77
India	1.47	31.84	Thailand	0.87	55.71
Indonesia	1.35	33.87	Trinidad	0.74	41.4

between the income distribution and the price of nontradables (proxied by the consumption price level from the Penn World Table 6.3) and the price of nontradables and real exchange rates. All data are transformed into logarithms. Table 4 reports these results.

Panel A of Table 4 reports the correlation analysis between $GINI$ and the price of nontradables. When all 69 countries are included, the correlation coefficient is positive (0.2097) and significant at the 5% critical

TABLE 1—*Continued*

Country	Real Exchange Rate REX	Gini Coefficients $GINI$	Country	Real Exchange Rate REX	Gini Coefficients $GINI$
Iran	1.42	41.18	Tunisia	1.98	41.33
Ireland	0.65	31.59	Uganda	3.72	42.35
Italy	0.73	31.57	U.K	0.9	31.2
Jamaica	0.78	40.96	U.S.A	0.81	43.62
Japan	0.74	28.75	Venezuela	1.21	45.48
Jordan	1.41	38.43	Zimbabwe	1.17	73.1
Korea, R.	0.69	33.73			

TABLE 2.

Mean values of income share of the lowest ($LQ1$) and highest ($LQ5$) quintiles, 1980-2007

Country	Income Share Lowest Quintile $LQ1$	Income Share Highest Quintile $LQ5$	Country	Income Share Lowest Quintile $LQ1$	Income Share Highest Quintile $LQ5$
Algeria	NA	NA	Lesotho	NA	NA
Australia	0.045167	0.463167	Malawi	0.050658	0.561357
Austria	0.0675	0.3875	Malaysia	0.0458	0.5373
Bangladesh	0.088791	0.419139	Mauritania	0.062	0.457
Belgium	NA	NA	Mauritius	0.0707	0.3992
Bolivia	0.04	0.491	Mexico	0.041	0.553
Brazil	0.03085	0.60005	Morocco	0.065	0.466
Cameroon	0.056	0.508	Nepal	0.075647	0.470293
Canada	0.0684	0.37722	Netherlands	NA	NA
Chile	0.039	0.5935	New Zealand	0.056663	0.404825
China	0.075708	0.380817	Nigeria	0.050461	0.491722
Colombia	0.037	0.559	Norway	NA	NA
Costa Rica	0.044	0.486	Pakistan	0.079986	0.4305
Cote d'Ivoire	0.053528	0.509086	Panama	0.0419	0.5242

level. We also test country-specific correlations for nine countries whose Gini coefficients are available for more than 10 years. While correlation coefficients are insignificant for Canada, Italy, Poland, and Sweden, other countries with significant correlation coefficients (with the exception of the Netherlands) have the expected positive sign. This means that when the income distribution deteriorates (a higher Gini), the price of nontradables

TABLE 2—*Continued*

Country	Income Share Lowest Quintile $LQ1$	Income Share Highest Quintile $LQ5$	Country	Income Share Lowest Quintile $LQ1$	Income Share Highest Quintile $LQ5$
Denmark	0.0667	0.3721	Peru	NA	NA
Dom. Rep.	0.041	0.5455	Philippines	0.055647	0.519837
Fiji	NA	NA	Poland	0.10598	0.35291
Finland	NA	NA	Portugal	0.0614	0.4042
France	NA	NA	Rwanda	0.05	0.52
Germany	NA	NA	Singapore	0.0376	0.485557
Ghana	NA	NA	South Africa	0.034745	0.621776
Greece	0.067267	0.3829	Spain	0.0897	0.345783
Guatemala	NA	NA	Sri Lanka	0.070066	0.473583
Honduras	0.0384	0.5633	Sweden	0.075667	0.368
Hong Kong	0.052733	0.4967	Switzerland	NA	NA
Hungary	0.0865	0.397	Taiwan	0.076533	0.3865
India	0.087867	0.411378	Thailand	0.052494	0.513838
Indonesia	0.077619	0.433824	Trinidad	0.0343	0.4486
Iran	0.057786	0.474981	Tunisia	0.06	0.473
Ireland	NA	NA	Uganda	0.057001	0.525084
Italy	0.073125	0.39975	U.K	NA	NA
Jamaica	0.060747	0.475535	U.S.A	0.038778	0.465
Japan	0.061467	0.406633	Venezuela	NA	NA
Jordan	0.066	0.4205	Zimbabwe	NA	NA
Republic of Korea	0.075	0.393			

goes up. This is consistent with our first assumption — that the higher income group has greater demand elasticity for the nontradables.

Panel B of Table 4 reports the correlation analysis between the price of nontradables and the real exchange rates. The correlation coefficient when all 69 countries are analyzed is -0.3188, which is significant at the 1% critical level. For the country-specific analysis, six out of the eight countries have the expected negative sign and their correlation coefficients are significant at the 1% critical level (with the exception of the Netherlands with significance at the 5% critical level).

These simple correlation analyses support our assumptions and proposed linkages — that changes in the income distribution affect the real exchange rate through a change in the price of nontradables.

TABLE 3.

Descriptive Statistics for Sample of 69 Countries over the Period 1980-2007

Panel A: Full Sample

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
REX	1932	1.300615	0.908055	0.1769	11.4832
$GINI$	585	38.10293	10.02456	20.324	73.1
$RCGDP$	1932	9944.849	9513.811	385.7999	53967.52
TOT	1798	1.049431	0.269067	0.3974	3.1563
$OPEN$	1871	0.586909	0.474775	0.0397	3.8596
LLY	1763	0.867236	8.797938	0.0571	289.1615
$FXKY$	1874	0.297517	0.594333	0.0353	5.7433
SCH	1276	12.0334	3.3577	4	20.7

Panel B: OECD Countries (25 countries)

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
REX	700	0.811905	0.142956	0.563	1.5435
$GINI$	318	32.30104	6.146882	20.324	54.6545
$RCGDP$	700	18451.35	8760.523	2687.648	53967.52
TOT	662	1.013713	0.179214	0.5885	3.0656
$OPEN$	687	0.525905	0.267249	0.1345	1.8635
LLY	591	0.716516	0.379033	0.1104	2.4385
$FXKY$	700	0.412044	0.949217	0.1408	5.7433
SCH	617	14.607	1.9971	10	20.7

Panel C: Non-OECD Countries (44 countries)

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
REX	1232	1.578291	1.033858	0.1769	11.4832
$GINI$	267	45.01305	9.350314	22.369	73.1
$RCGDP$	1232	5111.608	5820.325	385.7999	48489.63
TOT	1136	1.070246	0.3078	0.3974	3.1563
$OPEN$	1184	0.622306	0.558105	0.0397	3.8596
LLY	1172	0.94324	10.78793	0.0571	289.1615
$FXKY$	1174	0.22923	0.120877	0.0353	1.4235
SCH	659	9.62382	2.46718	4	14.6

6. FIXED — AND RANDOM — EFFECTS ESTIMATION MODELS

We investigate random- and fixed-effects models in this section; in all cases the dependent variable is the (log) real exchange rate, REX, and independent variables are also in logs. Table 5 reports estimated reduced forms where random-effects models can be rejected based on the Hausman test statistic. The different columns are for different country groupings; the first column is for the Latin American countries, whereas East Asian

TABLE 4.

Correlation Analyses

Panel A: Correlation between log(*Gini*) and log(*Price of nontradables*)

Sample	Correlation Coefficient	p-VALUE
All 69 countries	0.2097*	0.042
	Degrees of Freedom = 489	
Australia	0.9386**	0.0000
Canada	0.1744	0.4496
Italy	−0.0672	0.8537
Netherlands	−0.7587**	0.0026
New Zealand	0.7630*	0.0276
Poland	0.9217	0.0001
Sweden	0.2875	0.4531
United Kingdom	0.9273**	0.0000
United States	0.9732**	0.0000

Panel B: Correlation between log(*Real exchange rates*) and log(*Price of nontradables*)

Sample	Correlation Coefficient	p-VALUE
All 69 countries	−0.3188**	0.0000
	Degrees of Freedom = 1835	
Canada	−0.1671	0.3956
Italy	−0.5613**	0.0019
Netherlands	−0.4184*	0.0267
New Zealand	−0.5792**	0.0012
Poland	−0.6459**	0.0002
Sweden	−0.0846	0.6688
United Kingdom	−0.5688**	0.0016
United States	−0.9723**	0.0000

Note: Double asterisks (**) imply that the coefficient is significant at the 1% critical level and a single asterisk (*) implies that the coefficient is significant at the 5% critical level. Note also that Panel B has many more observations than does Panel A because there are many missing values for GINI.

countries are in the second column. The estimated $GINI$ coefficients are significant at the 1% critical level and negative, which is consistent with our expectation from Equation (3). This implies that an improvement in a country's income distribution may be associated with a devaluation of the domestic currency. As can be seen from both columns, this result of a significant, negative relationship between Gini coefficients and real exchange rates holds for the different groups of countries.

With the exception of LLY (liquid liability adjusted by GDP), other explanatory variables in Table 5 are also significant and most of them have the expected signs. $FXKY$ (public investment on fixed capital divided by GDP) is significant and negative. SCH (human capital) is positive and

TABLE 5.

Panel regression: Fixed-effects model

Dependent variable log(REX) & income distribution log($GINI$)

Independent Variables	(1)		(2)	
log($GINI$)	-0.226^{**}	(-2.865)	-0.250^{**}	(-2.778)
log(LLY)	0.032	(1.203)	0.036	(1.319)
log($FXKY$)	-0.153^{**}	(-5.459)	-0.151^{**}	(-5.319)
log($OPEN$)	0.082^{*}	(2.383)	0.073^{*}	(2.165)
log(SCH)	0.221^{**}	(2.692)	0.217^{**}	(2.659)
log($GINI$) $*$ $LATIN$	0.284	(0.928)	-	-
log($GINI$) $*$ $EASIA$	-	-	0.132	(0.919)
Adj. R-sq.	0.964		0.964	
Degrees of Freedom	230		230	
Hausman Test (p-value)	0		0	

Note: t-statistics are in parentheses; Double asterisks (**) imply that the coefficient is significant at the 1 % critical level and a single asterisk (*) implies that the coefficient is significant at the 5 % critical level. We can reject the random-effects model from the significance probability of Hausman test statistics.

significant. However, the measure of openness, $OPEN$, has an unexpected positive sign.

Another important question is whether countries with different income levels have different transmission mechanisms from income inequality to the real exchange rate. We use interaction terms of dummy variables ($OECD$, East Asian ($EASIA$) and Latin America ($LATIN$)) and log($GINI$) in the regression to capture the possible role of income levels (or regional differences) on the transmission mechanism. However, Table 5 shows that the interaction terms of dummy variables and log($GINI$) are insignificant for all cases.

The findings of the fixed-effect models in Table 5 support the proposition presented above, that an improvement in the income distribution is associated with real exchange-rate depreciation. This finding also has an important policy implication: that reducing inequality, by decreasing the price of nontradables, will depreciate the real exchange rate, which may well help to increase the international competitiveness of the tradable sector. Because an appreciation of the real exchange rate might hurt an economy's international competitiveness, income inequality could have a negative implication for sound macroeconomic management. However, the robustly negative relationship between the real exchange rate and income inequality does not imply that dramatic redistributive policies will automatically bring about a real depreciation of the domestic currency, thereby improving the external balance and accelerating economic growth. Overly ambitious income redistributive policies might cause domestic inflation and could distort incentives and hurt productivity, all of which would slow down long-term economic growth (see, e.g., Al-Marhubi, 1997).

7. SENSITIVITY ANALYSIS

Supplementary regressions for Equation (9) are run using income shares of the lowest and highest quintiles, $LQ1$ and $LQ5$, respectively, as measures of income inequality. These estimation results are reported in Tables 6 and 7.

Table 6 reports estimation results of the random-effects model when inequality is measured by the income share of lowest quintile ($LQ1$) of population. The positive sign on the lowest-income-quartile variable indicates that an increase in the income share of the lowest quintile (that is, a decrease in inequality) will bring about a real depreciation of the exchange rates (significance is at the 1% critical level). Increasing the income share of the lowest quartile will decrease the demand for nontradables and thus reduce their price. This finding is consistent with our previous findings in which the Gini was used as an income inequality measure.

TABLE 6.

Panel regression: Random-effects model

Dependent variable log(REX) & income distribution lowest quintile income share log($LQ1$)

Independent Variables	(1)		(2)	
Constant	-0.375	(-1.177)	-1.122^{**}	(-2.844)
log($LQ1$)	0.205^{**}	(3.141)	0.180^{**}	(2.825)
log(LLY)	0.159^{**}	(3.421)	0.086	(1.676)
log($OPEN$)	0.074	(1.179)	0.13^{*}	(2.08)
log($RCGDP$)	-	-	-0.185^{**}	(-3.084)
log(SCH)	0.269^{*}	(2.093)	0.497^{**}	(3.45)
Log Likelihood	38.981		42.867	
Degrees of Freedom	122		121	
Hausman Test (p-value)	0.076		0.290	

Note: t-statistics are in parentheses; Double asterisks (**) imply that the coefficient is significant at the 1% critical level while single asterisk (*) implies that the coefficient is significant at the 5% critical level. We cannot reject the random-effects model from the significance probability of Hausman test statistics.

Table 7 shows the estimation results of the reduced-form model when the income share of the highest quintile ($LQ5$) is used as a measure of income inequality. The estimated coefficients of the highest income quintile variable ($LQ5$) indicate that an increase in the income share of the highest quintile (that is, an increase in inequality) will bring about a real appreciation of the exchange rates (significance is at the 1% critical level).

Increasing the income share of the highest quintile will increase the demand for nontradables and thus increase their price. This finding is consistent with our previous findings in which the Gini was used as an income inequality measure. The estimated coefficients for $FXKY$ and TOT are significant and have the expected negative signs.

TABLE 7.

Panel regression: Fixed-effects model

Dependent variable log(REX) & income distribution highest quintile income share log($LQ5$)

Independent Variables	(1)		(2)	
log($LQ5$)	-0.530^{**}	(-3.398)	-0.579^{**}	(-3.713)
log(LLY)	0.054	(0.999)	0.051	(0.924)
log($FXKY$)	-0.032^{**}	(-2.956)	-0.024^{*}	(-2.425)
log($OPEN$)	0.068	(0.707)	0.138	(1.552)
log(SCH)	0.009	(0.036)	0.056	(0.212)
log($RCGDP$)	0.208	(1.713)	-	-
log(TOT)	-0.271^{**}	(-2.800)	-0.273^{**}	(-2.771)
Adj. R-sq.	0.968		0.967	
Degrees of Freedom	51		52	
Hausman Test (p-value)	0.005		0.005	

Note: t-statistics are in parentheses; Double asterisks (**) imply that the coefficient is significant at the 1% critical level and a single asterisk (*) implies that the coefficient is significant at the 5% critical level. We can reject the random-effects model from the significance probability of Hausman test statistics.

8. DYNAMIC PANEL ESTIMATION

Because there is a high degree of persistence (or inertia) in real exchange rates, REX, we estimate a dynamic model that includes the lagged dependent variable, the log of the real exchange rate:

$$\begin{aligned} REX_{i,t} &= a_{i,t} + bREX_{i,t-1} + c[Income\ distribution]_{i,t} \\ &+ d[Control\ variables]_{i,t} + \eta_i + \nu_{it} \end{aligned} \tag{10}$$

where η_i is an unobserved country-specific effect and ν_{it} is a disturbance term. Because the explanatory variables and dependent variable may be correlated with η_i a transformation such as first differencing is required. Following Maudos and Solis (2009, p. 1928), we use a methodology proposed by Arellano and Bover (1995) and Blundell and Bond (1998) to estimate a system of equations in both first-differences and levels, the "system" GMM estimator that combines the set of equations in first-differences (with suitably lagged levels as instruments) with another set of equations in levels (with suitably lagged first-differences as instruments). In Table 8, two-step GMM estimators are used with asymptotic standard errors that are robust to heteroscedasticity.

TABLE 8.

System-GMM estimation

Dependent variable $\log(REX)$ & income distribution $\log(GINI)$

Independent Variables	(1)		(2)	
$\log(REX)_{t-1}$	0.666^{**}	(43.710)	0.725^{**}	(41.541)
$\log(GINI)$	-0.045^{**}	(-3.811)	-0.071^{**}	(-9.612)
$\log(LLY)$	0.039^{**}	(4.920)	0.031^{*}	(2.091)
$\log(OPEN)$	-0.025^{**}	(-5.540)	-0.029^{**}	(-9.019)
$\log(SCH)$	-0.006	(-0.488)	-0.041	(-1.256)
$\log(RCGDP)$	-0.078^{**}	(-11.983)	-0.049^{**}	(-5.031)
$EASIA$	-0.256	(-0.704)	-	-
$\log(GINI) * EASIA$	-0.196	(-0.545)	-	-
$LATIN$	-	-	0.089	(1.872)
$\log(GINI) * LATIN$	-	-	0.090	(1.085)
Arellano-Bond test for AR(1)	-0.84	[0.401]	-0.92	[0.358]
Arellano-Bond test for AR(2)	-0.69	[0.492]	-0.69	[0.488]
Hansen J Test	21.22	[0.507]	23.83	[0.413]
Number of Instruments	31		32	
Number of Observations	108		108	

Note: t-statistics are in parentheses; double asterisks (**) imply that the coefficient is significant at the 1% critical level while a single asterisk (*) implies significance at the 5% critical level; figures in brackets are significance probability; Arellano-Bond test for AR1 (or 2) are tests for first- (or second-) order serial correlation, respectively; the Hansen J test is a test of over-identification restrictions.

Column (1) of Table 8 shows the system-GMM estimation results for Equation (10) with $EASIA$ country dummies and the interaction term $\log(GINI)$ and $EASIA$ country dummy variables; whereas column (2) shows the system-GMM estimation results with Latin American dummies and the interaction term $\log(GINI)$ and $LATIN$ regional dummy variables. Using Hansen's test of over-identifying restrictions we could not reject the null hypothesis that the model is correctly specified and the instruments are valid. We also report Arellano and Bond's (1991) tests for autocorrelations of order 1 and 2; these test results consistently show that we cannot reject the null hypothesis of no second-order autocorrelations.

From Table 8, we can also see that the system-GMM-estimated coefficient for our inequality variable ($GINI$) has the expected negative sign and is significant at the 1% critical level. The significance of the lagged dependent variable, REX, implies that there is indeed inertia in the annual real exchange rate series we are using, thus validating the dynamic panel approach. While human capital (SCH) is insignificant, liquid liability (LLY) has the expected positive sign and openness ($OPEN$) has the expected negative sign. At the same time, real GDP per capita ($RCGDP$) is significant and has a negative sign, which is consistent with Kravis and Lipsey (1988). Overall, the system-GMM estimation results confirm the estimation results obtained using the fixed- and random-effects models. However, interaction terms of $\log(GINI)$ with dummy variables for Latin America and East Asia are insignificant.[16]

9. PANEL VECTOR AUTOREGRESSIONS

Finally, we use panel vector autoregression (VAR) analysis as it treats all the variables in the model as endogenous and allows for unobserved individual heterogeneity. The first-order panel VAR model is specified as follows:

$$z_{i,t} = \Gamma_0 + \Gamma_1 z_{i,t-1} + f_i + d_{c,t} + e_t \tag{11}$$

where z is either a four-variable vector $\{REX, GINI, RCGDP, OPEN\}$ or a five-variable vector $\{REX, GINI, RCGDP, OPEN, LLY\}$ and $\{REX, GINI, RCGDP, OPEN, FXKY\}$. All variables are defined as they were above.

Our main objective is to compare the response of real exchange rates to income distribution across countries. Table 9 reports the estimated coefficients for the panel VAR with four or five variables. Panel A reports the results for $\{REX, GINI, RCGDP, OPEN\}$; Panel B reports the results for $\{REX, GINI, RCGDP, OPEN, LLY\}$; and Panel C reports the results for $\{REX, GINI, RCGDP, OPEN, FXKY\}$. All three panels support our hypothesis as they show that the response of real exchange rates to Gini coefficients is always negative and significant at the 1% critical level.

10. CONCLUSIONS

This paper provides robust empirical evidence for a negative association between income inequality and the real exchange rate. First, we showed that income inequality is positively related with the price of nontradables. Second, we demonstrated that an improvement in the income distribution, through a decline in the price of nontradables, will depreciate the real exchange rate. The magnitude of the association between various measures of income inequality and the real exchange rate is large and the estimation results are robust to alternative specifications of the reduced-form equations

[16]Due to insufficient data, we could not perform system-GMM estimation using $LQ1$ and $LQ5$ (lowest and highest income share quintiles, respectively) as proxies for income inequality.

TABLE 9.

Estimation results of the panel vector autoregressions (full sample)

Panel A: $\{REX, GINI, RCGDP, OPEN\}$

Response of	Response to			
	$REX(t-1)$	$GINI(t-1)$	$RCGDP(t-1)$	$OPEN(t-1)$
$REX(t)$	0.569 (2.08)**	−0.112 (−8.55)***	7.81E-06 (1.95)*	−0.029 (−0.13)
$GINI(t)$	−2.614 (−1.02)	0.430 (4.31)***	1.92E-06 (0.05)	2.681 (1.43)
$RCGDP(t)$	1531 (1.14)	−89.79 (−1.92)*	1.025 (45.85)***	−17.62 (−0.02)
$OPEN(t)$	0.563 (2.24)**	−0.012 (−1.12)	9.52E-06 (2.18)**	0.606 (3.41)***
N obs	337			
N countries	69			

TABLE 9—*Continued*

Panel B: $\{REX, GINI, RCGDP, OPEN, LLY\}$

Response of	Response to				
	$REX(t-1)$	$GINI(t-1)$	$RCGDP(t-1)$	$OPEN(t-1)$	$LLY(t-1)$
$REX(t)$	−0.025 (−0.12)	−0.126 (−7.39)***	$-4.69E-06$ (−1.26)	0.091 (0.43)	0.600 (2.68)***
$GINI(t)$	−1.323 (−0.88)	0.503 (4.57)***	5.19E-05 (2.37)**	2.076 (1.59)	−2.283 (−1.66)*
$RCGDP(t)$	−2214 (−2.24)**	189.2 (1.98)**	1.071 (57.80)***	90.16 (0.07)	−3268 (−2.40)**
$OPEN(t)$	0.180 (1.66)*	−0.011 (−1.05)	2.75E-06 (1.73)*	0.714 (5.93)***	0.211 (1.66)*
$LLY(t)$	0.031 (0.78)	0.004 (1.21)	6.39E-07 (0.91)	0.072 (1.94)*	0.876 (19.89)***
N obs	271				
N countries	69				

Panel C: $\{REX, GINI, RCGDP, OPEN, FXKY\}$

Response of	Response to				
	$REX(t-1)$	$GINI(t-1)$	$RCGDP(t-1)$	$OPEN(t-1)$	$FXKY(t-1)$
$REX(t)$	0.660 (1.90)*	−0.112 (−7.78)***	8.52E-06 (1.73)*	−0.005 (−0.02)	−0.060 (−0.78)
$GINI(t)$	−3.291 (−1.05)	0.454 (3.97)***	$-4.91E-06$ (−0.10)	2.682 (1.31)	−0.357 (−0.59)
$RCGDP(t)$	1555 (0.87)	−89.19 (−1.59)	1.025 (38.01)***	15.05 (0.01)	−135.2 (−0.42)
$OPEN(t)$	0.747 (2.11)**	−0.016 (−1.17)	1.14E-05 (2.05)**	0.564 (2.66)***	0.112 (1.98)**
$FXKY(t)$	−0.122 (−1.17)	0.007 (2.04)**	$-2.83E-06$ (−1.75)*	0.148 (2.54)**	0.789 (7.88)***
N obs	335				
N countries	69				

Note: t-statistics are in parentheses; Double asterisks (**) imply that the coefficient is significant at the 1 critical level and a single asterisk (*) implies that the coefficient is significant at the 5% critical level.

and estimation methodologies. An important policy recommendation follows directly from our key finding: a sustainable redistributive policy that does not distort incentives may, by being associated with a real depreciation, accelerate the growth momentum of the economy. Finally, although the analysis has demonstrated a robust negative correlation between the real exchange rate and income inequality, the direction of causation has not been determined. It may also be desirable to look into a specific country's experience as a complement to this study.

APPENDIX: DEFINITIONS AND SOURCES OF VARIABLES

Dummy Variables: $EASIA$: 1 for East Asia, 0 otherwise; $LATIN$: 1 for Latin America, 0 otherwise; and $OECD$: 1 for member countries, 0 otherwise.

$FXKY$: Gross Fixed Capital formation to GDP. Gross fixed capital formation (IFS line 93.e) to GDP (IFS line 99.b) is measured by domestic public (and private) investment to GDP. This is used as a proxy for the manufacturing sector productivity. For Taiwan, data are from the National Statistics of Taiwan.

$GINI$: Gini coefficients are from UNU-WIDER World Income Inequality Database ($WIID$) V2.0c (2008). Given various sources of the coefficients, following criteria in the order are applied to sort out: data availability, date quality classification (provide by the source), data coverage (in terms of area, population, and age), etc.

LLY: Liquid liability to GDP is money plus quasi-money (IFS line 35.l) over GDP (IFS line 99.b). For Japan, money (IFS line 34) and quasi-money (IFS line 35) are separately extracted and then aggregated to derive liquid liability. For the euro-zone countries before the introduction of the euro, liquid liability is the sum of currency issued (IFS line 34.a), demand deposits (IFS line 34.b) and other deposits (IFS line 35). In Taiwan, M2, as a liquid liability and GDP are obtained from the Central Bank of Taiwan.

$LQ1$ and $LQ5$: The lowest quintile, $LQ1$, and highest quintile, $LQ5$, of the income share are from the UNU-WIDER World Income Inequality Database ($WIID$) V2.0c (2008).

$OPEN$: Trade Openness is measured as the sum of Exports (IFS line 70) and Imports (IFS line 71) divided by GDP (IFS line 99.b). For Mauritania, the sum of Goods: Exports (IFS line 78AA) Goods: Imports (IFS line 78AB) are counted, instead of Export and Import respectively. For Taiwan, these data are from the National Statistics of Taiwan.

$RCGDP$: Real GDP per capita is from the Penn World Table 6.3.

REX: Real Exchange Rate is defined as the relative price of tradables to nontradables. The price level of investment (pi) in Penn World Table 6.3 is used as the price of tradables, whereas the price level of consumption (pc) in Penn World Table 6.3 is used as the price of nontradables. The REX is defined as pi/pc.

SCH: Expected years in school, from primary to tertiary education (e.g., elementary school to college), or the average number of years that a child is likely to spend in the educational system of his or her country. UNESCO Institute for Statistics.

TOT: Terms of trade are calculated by dividing export price (IFS line 76) by import price (IFS line 76.x). For those countries with missing values in IFS, we use the net barter terms of trade (2000=1) in the World Bank's World Development Indicators. For Taiwan, these data are from the National Statistics of Taiwan.

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