Macroeconomic Dynamics, **20**, 2016, 1–26. Printed in the United States of America.
doi:10.1017/S1365100514000510

ARTICLES

GLOBALIZATION AND INFLATION: NEW PANEL EVIDENCE

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Recent empirical work on globalization and inflation analyzes multicountry data sets in panel and/or cross-section frameworks and reaches inconclusive results. This paper highlights their shortcomings and reexamines the issue utilizing heterogeneous panel cointegration techniques that allow for cross-section heterogeneity and dependence. It finds that in a sample of developing countries globalization of both trade and finance, on the average, exerts a significant and positive effect on inflation, whereas in a sample of developed countries there is, on the average, no significant impact of openness. Neither type of openness disciplines inflationary policy. Despite this, there are large variations in the effect across countries, due possibly to differences in the quality of political institutions, central bank independence, the exchange-rate regimes, financial development, and/or legal traditions.

Keywords: Trade Openness, Financial Openness, Inflation, Cross-Section Dependence, Heterogeneous Panels

1. INTRODUCTION

The past few decades have witnessed a trend toward worldwide integration, in terms of both trade and financial markets, along with a decline in the level of (expected) inflation in most advanced economies and some developing countries. This tendency for inflation to drop with greater globalization has been regarded as evidence of openness in disciplining monetary policy, one important collateral benefit of trade and financial account liberalization as put forward by Winters (2004) and Kose et al. (2009). Nonetheless, there is hardly any consensus at the theoretical level, as there are plausible models suggesting a disciplinary role of openness as well as models suggesting the opposite. Some simply regard

We are grateful to the associate editor and one anonymous referee for their valuable comments. This work was supported by Special Research Grant of Korea University, and by Taiwan's Ministry of Science and Technology (NSC 98-2410-H-032-31-MY2). The usual disclaimer applies. Address correspondence to: Shu-Chin Lin, Department of Economics, Tamkang University, 151 Yingzhuan Rd., Danshui Dis., New Taipei City 25137, Taiwan; e-mail address: e-mail: econscl@mail.tku.edu.tw.

 1365-1005/15

openness as irrelevant. On one hand, Romer (1993), Lane (1997), Rogoff (2003), and Daniels and VanHoose (2006), to name a few, argue that openness reduces output gains from unexpected monetary expansion and the incentives for monetary authorities to inflate, thereby constraining inflation. On the other hand, Cavelaars (2009), Cooke (2010), and Evans (2012) suggest the opposite. Rodrik (2001), Temple (2002), and Ball (2006) cast doubts on the disciplinary impact of openness.

A growing body of empirical literature has attempted to evaluate this debate but has reached inconclusive results. Some studies find a negative association between inflation and openness. Examples include Romer (1993), Lane (1997), Terra (1998), Gruben and McLeod (2004), Daniels et al. (2005), and Badinger (2009). Other studies such as Karras (1999) and Alfaro (2005) point into the direction of a positive effect. Works including Romer (1993), Cavallari (2001), and Badinger (2009) even find a nonlinear response of inflation to openness.

These empirical works analyze multicountry data sets in panel and/or cross-section frameworks and have several limitations. First, cross-country regressions of time averages emphasize the variation in the data across countries and therefore fail to address country-specific differences and dynamics in the inflation–openness relationship. Second, panel regressions focus on the within variation in the data and hence allow one to analyze how changes in openness relate to changes in inflation within a given country. However, they fail to address parameter heterogeneity and cross-section dependence. Countries differ in the extent and pace of globalization, in the formation and implementation of inflation management policy, and in other political and social characteristics. These structural makeups tend to be rigid, requiring significant time and effort for any change, which may cause differences in the sensitivity of inflation to openness across countries. Lack of correspondence between panel and country-specific estimates may make the generalizations based on panel estimates, i.e., the very "broad conclusion," proffer incorrect inferences with limited policy relevance. Pesaran et al. (2000) point out the issue of parameter heterogeneity across panel units (countries) and show that unless this issue is addressed, panel estimates become biased and inconsistent. On the other hand, cross-section dependence may arise because of spatial correlations, spillover effects, omitted global variables, and common unobserved shocks. Cross-section correlation can potentially induce serious bias into the estimates because the impact assigned to an observed covariate in reality confounds its impact with that of the unobserved processes [Pesaran (2006)]. Third, many macroeconomic variables are potentially nonstationary [Nelson and Plosser (1982); Granger (1997)], a property that cannot be rejected for the variables in our data. When variables are nonstationary, averaging the data, for example, over five years or longer, cannot solve the problem of unit roots in the data because the data-generating process remains unchanged. There is a risk of performing spurious regression in the presence of nonstationary variables.

Last, the implications of globalization for inflation have been studied primarily in relation to the degree of exposure to international trade. The evidence on the effects of financial openness is quite sparse.[1] As pointed out by Badinger

(2009), this gap in the literature is particularly surprising in light of the fact that financial account openness has increased dramatically since the 1990s and that a comprehensive, high-quality data set for financial openness has been developed recently by Lane and Milesi-Ferretti (2007). Furthermore, there is a growing literature pointing to substitutability or complementarity between trade and financial openness [Aizenman (2004, 2008); Aizenman and Noy (2009)]. From an econometric perspective, excluding a trade or financial openness variable from an analysis that considers the relationship between inflation and openness can thus cause bias in the results.

This paper reconsiders the openness–inflation nexus. Particularly, it applies heterogeneous panel cointegration techniques that allow one to address endogeneity of the regressors, parameter heterogeneity, cross-section dependence, and nonstationarity in a common factor-modeling framework [Pesaran (2006); Bai (2009); Kapetanios et al. (2011)]. Hence, apart from recognizing the potential for differences and commonalities among countries, the methodology considers the individual evolution of countries over time. The analysis represents a step toward making cross-country empirics relevant to individual countries by moving away from empirical results that characterize the average country and toward a deeper understanding of the differences.

To anticipate the results, we find that during the period 1970–2007, globalization of both trade and finance, on the average, exerts a significant and positive effect on inflation in a sample of developing countries. As for advanced countries, there is, in general, no significant impact of openness. The data suggest that openness does not constrain inflation in either the developed or the developing countries. However, the result for the whole sample does not imply that both types of openness do not discipline monetary policy in a country. Particularly, our cross-country regression results suggest that the differences in the inflation effect of openness among countries depend upon differences in the quality of political institutions, central bank independence, the exchange-rate regimes, financial development, and/or legal traditions.

The remainder of the paper is organized as follows. Section 2 introduces the model specification and the estimation methodology and describes the data. Section 3 analyzes the empirical results. Section 4 explains differential long-run impacts of openness on inflation. In Section 5, we conclude the analysis.

2. METHODOLOGY AND DATA

2.1. Methodology

To evaluate whether trade and financial openness contribute to differences in inflation across countries over time, we consider the following common-factor regression model:

$$y_{it} = \beta_i' x_{it} + u_{it}, \tag{1}$$

$$u_{it} = \alpha_i + \lambda_i' f_t + \varepsilon_{it}, \tag{2}$$

https://doi.org/10.1017/S1365100514000510 Published online by Cambridge University Press

$$x_{mit} = \pi_{mi} + \rho'_{mi} f_{mt} + \gamma'_{mi} g_{mt} + e_{mit}, \quad (3)$$

$$f_t = \phi'_i f_{t-1} + \nu_{it}, \quad g_t = \theta'_i g_{t-1} + \omega_{it}, \quad (4)$$

where $i = 1, 2, \ldots, N$ is the country indicator, $t = 1, 2, \ldots, T$ is the time index, $m = 1, 2, \ldots, K$ is the number of observed regressors in x, including trade and financial openness indicators, and $f_{mt} \subset f_t$. y is an inflation indicator. β'_i is a set of the country-specific slopes on the observable regressors. u_{it} contains the unobservable terms and the white noise error terms ε_{it}. The unobservables in equation (2) are made up of the group fixed effects α_i, which capture time-invariant heterogeneity across groups, as well as an unobservable common factor f_t with heterogeneous factor loadings λ'_i, which can capture time-variant heterogeneity and cross-section dependence. Likewise, in equation (3), observable $x's$ are modeled as linear functions of the white noise error terms e_{mit} and the unobserved common factors f_{mt} and g_{mt}, with respective country-specific factor loadings ρ'_{mi} and γ'_{mi}, which capture time-variant heterogeneity and cross-section dependence. $x's$ are driven by f_{mt}, a subset of the factor driving y; this leads to endogeneity, whereby the regressors are correlated with the unobservable, making it difficult to identify β'_i separately from ρ'_{mi} and γ'_{mi} [Kapetanios et al. (2011)]. Equation (4) indicates that the unobservable common factors follow an AR(1) process with white noise error terms ν_{it} and ω_{it}. This allows for nonstationarity in the factors if $\phi = 1$ and $\theta = 1$ and hence in the observables.

The nature of macroeconomic variables in an integrated world, where economies are strongly connected to each other and latent forces drive all of the outcomes, provides a conceptual justification for the pervasive character of unobserved common factors. However, the presence of these latent factors makes it difficult to argue for the validity of traditional approaches to causal interpretation of cross-country empirical analyses. Instrumental variable estimation in cross-section regressions or Arellano and Bond-type (1991) lag instrumentation within pooled-panel models becomes invalid in the face of common factors and/or heterogeneous equilibrium relationships [Pesaran and Smith (1995)]. Alternatively, the model can be estimated using a heterogeneous (between dimensions) estimator based on the mean group (MG) estimator of Pesaran and Smith (1995). The MG accounts for heterogeneous panels by running separate ordinary least squares (OLS) regressions to obtain the individual slope estimates of β_i for each cross-section unit. The estimated coefficients $\hat{\beta}_i$ are subsequently averaged across each cross section. However, this estimator has little concern with cross-section dependence and assumes away $\lambda'_i f_t$ or at best models these unobservables with a linear trend. Coakley et al. (2006) show, for nonstationary and cross-section-dependent data, that MG estimates are severely affected by their failure to account for cross-section dependence.

To correct the drawback, the Common Correlated Effects (CCE) estimator developed by Pesaran (2006) and extended by Kapetanios et al. (2011) and Pesaran and Tosetti (2011) accounts for the presence of unobserved common factors with heterogeneous impacts by including cross-section averages of the dependent

($\bar{y}_t$) and independent variables ($\bar{x}_t$) in the regression. To obtain insight into the mechanics of this approach, consider the cross-section average of equations (1) and (2). As the cross-section dimension N increases, given $\bar{\varepsilon}_t = 0$, we obtain

$$\bar{y}_t = \bar{\alpha} + \bar{\beta}'\bar{x}_t + \bar{\lambda}' f_t \quad \Leftrightarrow \quad f_t = \bar{\lambda}^{-1}(\bar{y}_t - \bar{\alpha} - \bar{\beta}'\bar{x}_t), \tag{5}$$

where $\bar{y}_t = N^{-1}\sum_{i=1}^{N} y_{it}$, and $\bar{x}_t = N^{-1}\sum_{i=1}^{N} x_{it}$. The unobserved common factor f_t can be captured by a combination of cross-sectional averages of y and x, i.e., $\bar{y}_t$ and $\bar{x}_t$. Modifying equations (1) and (2) accordingly, we have

$$y_{it} = \alpha_i + \beta_i' x_{it} + c_{0i}\bar{y}_t + c_i'\bar{x}_t + v_{it}. \tag{6}$$

MG estimation of (6) provides consistent estimates of the model parameters as simple averages of the country-specific estimates, e.g.,

$$\hat{\beta}_{CCEMG} = N^{-1}\sum_{i=1}^{N} \hat{\beta}_i.$$

This is the MG version (CCEMG) of the CCE estimator. In the pooled version (CCEP) of CCE estimators, the cross-section averages are interacted with country dummies, so that each country can have a different parameter on the cross-section averages:

$$y_{it} = \alpha_i + \beta' x_{it} + \sum_{j=1}^{N} c_{0i}(\bar{y}_t D_j) + \sum_{j=1}^{N} c_i'(\bar{x}_t D_j) + v_{it}, \tag{7}$$

where D_j represents country dummies. The CCEMG is thus a simple extension to the MG estimator based on country-specific OLS regressions, whereas the CCEP is a standard fixed effects (FE) estimator augmented with additional regression terms.

It has been shown that the CCE estimators are able to accommodate the type of endogeneity introduced into the regression equation and to yield consistent estimates for common β coefficients or the means of heterogeneous β_i [Coakley et al. (2006); Pesaran (2006); Kapetanios et al. (2011)]. This result is robust even when the cross-section dimension N is small, when variables are nonstationary, cointegrated, or subject to structural breaks, and/or when there are local spillovers and global/local business cycles [Chudik et al. (2011); Kapetanios et al. (2011); Pesaran and Tosetti (2011)]. Thus, we can exploit all the information available in the data set using annual-data estimation without incurring the distorting influence normally associated with business cycle components in this type of empirical analysis [Eberhardt and Teal (2011, 2013)].

2.2. Data

Our sample is an annual and unbalanced panel over the period from 1970 to 2007, consisting of 83 developing countries and 39 developed countries. Countries are selected if we have consecutive observations over at least half of the sample period.

TABLE 1. Descriptive statistics, cross-section-independence, and panel unit root tests

	inf_cpi	*inf_gdp*	*trade*	*fopen*
	Panel A: Descriptive statistics			
Mean	4.7348	4.7377	4.1126	4.5878
Std.	0.2811	0.2916	0.6545	0.8678
Min.	4.3609	4.2602	−1.1751	1.7113
Max.	9.3801	9.4286	6.0808	10.1015
Obs.(N*T)	3,927	4,184	4,187	4,194
	Panel B: Cross-section-independence and panel unit root tests			
	Developing countries			
CD test statistics	64.7520***	48.0210***	66.9280***	176.6080***
IPS: levels	−18.1312***	−13.7504***	−5.1895***	0.5943
First differenes	−34.3937***	−41.2343***	−27.6764***	−22.9562***
CIPS: levels	−2.6330	−2.7810	−2.2050	−2.6800
First differenes	−5.3110***	−6.3220***	−4.5560***	−4.2740**
Obs.	2,599(81)	2,845(83)	2,833(83)	2,858(83)
	Advanced countries			
CD test statistics	75.2090***	66.8230***	61.2570***	119.2420***
IPS: levels	−6.7233***	−9.1606***	−1.1980	2.5243
First differenes	−22.9051***	−34.4090***	−24.0270***	−23.7171***
CIPS: levels	−6.0570***	−3.2510	−2.2020	−2.3790
First differenes	−5.7450***	−7.1080***	−5.2600***	−5.1450***
Obs.	1,328(39)	1,339(39)	1,354(39)	1,336(39)

Note: For the CD test, each variable is regressed on a country-specific intercept. The IPS and $\Delta q_{it} = a_i + b_i t + c_i q_{i,t-1} + d_i \Delta q_{i,t-1} + \varepsilon_{it}$. The CIPS panel unit root test is based on the ADF regression: $\Delta q_{it} = a_i + b_i t + c_i q_{i,t-1} + d_i \Delta q_{i,t-1} + g_i \bar{z}_{t-1} + \varepsilon_{it}$, where $\bar{z}_t = (\bar{q}_{t-1}, \Delta \bar{q}_t, \Delta \bar{q}_{t-1})'$. The number of lags was determined by the Schwarz criterion with a maximum of five lags.
*** Significant at the 1% level.

As in Romer (1993), our dependent variable is the natural logarithm of the inflation rates, calculated with the GDP deflator (base year 2000). As a robustness check, we experiment with CPI-based inflation. In measuring trade and financial openness, we use de facto measures. As Temple (2002) notes, the theoretical predictions are based on the importance of trade relative to GDP, not trade policy. The same argument applies to financial openness. We then follow the common practice of using the trade share measure (*trade*), the (logarithm of) sum of exports and imports as a percentage of GDP, as our preferred indicator.[2] This measures actual exposure to trade interactions and accounts for the effective level of integration. Similarly, we consider a de facto measure of financial openness (*fopen*), the sum of foreign assets and liabilities as a share of GDP, which we calculate from the data set developed by Lane and Milesi-Ferretti (2007).[3] This measure is the outcome of the interaction between market forces and the enforcement of existing regulation. Table 1 reports the summary statistics (Panel A) and time-series properties of variables (Panel B).

As indicated in Panel B, our sample is characterized by cross-section dependence. The cross-section dependence (CD) test of Pesaran (2004) rejects the null hypothesis of no cross-section dependence for all variables. Also in Panel B, although the Im et al. (2003) panel unit-root test (IPS) shows stationarity in variables except for the trade or financial openness variable, the cross-section-dependent panel unit-root test (CIPS) of Pesaran (2007) suggests that most variables are nonstationary, i.e., integrated of order one, $I(1)$. Because the standard panel unit-root tests will show large size distortions when the data contain cross-section dependence across countries, and sizeable amounts of cross-section dependence are detected, we conclude that most variables considered are nonstationary, except for the CPI-based inflation in developed countries. Appendix A gives a brief introduction to the CIPS and CD tests.

3. EMPIRICAL RESULTS

In this section, we first examine whether and how trade and financial openness affect inflation in the long run. We then check whether the results are robust to alternative sample periods and whether there are significant differences in the long-run effects of openness on inflation across countries. Finally, we look at whether and how both types of openness are associated with inflation in the short run through panel error-correction models.

3.1. Long-Run Estimates

Table 2 reports the estimation results from the heterogeneous models for the developing and developed country samples. The estimation uses either GDP-deflator-based or CPI-based inflation. As a comparison, the pooled estimates are reported in Table B.1 of Appendix B. Specifically, in the heterogeneous models, we implement the MG and CCEMG estimators, both of which allow for heterogeneous slopes (β_i) but which differ in their assumptions about common factors. In the pooled models we implement the FE and CCEP estimator; both assume common slope parameters (β), but again differ in their assumptions about common factors. For all regression models, we report residual diagnostic tests, including Pesaran's CIPS and CD tests, which we use to build our judgment for a preferred empirical model. Residual nonstationarity invalidates the inferential tools employed (for instance, t-statistics) and indicates that regression results are potentially spurious, in the same way that residual cross-section dependence violates the assumption that the error terms are independently and identically distributed. This suggests that the specific model tested fails to adequately address the cross-country correlation of openness, inflation, and the unobservable induced by, for instance, common shocks or local spillover effects.[4]

For the full sample, Panel A of Table 2 presents the robust means for each regressor across N country regressions. In the developing country sample, MG estimates of the financial openness variables are negative but not statistically significant,

TABLE 2. Estimates of the long-run effect of openness on inflation

	Developing countries				Advanced countries			
	MG		CCEMG		MG		CCEMG	
	inf_gdp	*inf_cpi*	*inf_gdp*	*inf_cpi*	*inf_gdp*	*inf_cpi*	*inf_gdp*	*inf_cpi*
Panel A: 1970–2007								
trade	0.0897	0.1632*	0.1108*	0.1963***	0.1087*	0.0590	0.0172	−0.0002
	(0.0844)	(0.0859)	(0.0636)	(0.0674)	(0.0654)	(0.0615)	(0.0430)	(0.0236)
fopen	−0.0589	−0.0779	0.1229**	0.0948*	−0.0622***	−0.0263	−0.0236	−0.0007
	(0.0410)	(0.0633)	(0.0613)	(0.0540)	(0.0231)	(0.0408)	(0.0281)	(0.0238)
Diagnostics								
CIPS test	−4.1730***	−3.9790**	−6.0410***	−5.6850***	−3.5730**	−3.1710*	−4.5290***	−4.1060***
ΔCIPS	−7.7780***	−6.7670***	−8.6530***	−7.9200***	−6.6690***	5.8920***	−7.5060***	−6.5790***
CD tests	31.5090***	34.9800***	1.4490	1.3190	20.7880***	24.5320***	1.6180	1.1380
Panel B: 1970–1989								
trade	0.0375	0.1247**	−0.0723	−0.0002	0.0319	0.0875**	−0.1740**	−0.1132**
	(0.0669)	(0.0630)	(0.0476)	(0.0605)	(0.0565)	(0.0339)	(0.0762)	(0.0506)
fopen	0.0960*	0.0372	0.1119	0.1227	−0.0301	−0.0126	0.0583	0.0462
	(0.0494)	(0.0692)	(0.0789)	(0.0797)	(0.0406)	(0.0397)	(0.0414)	(0.0367)
Diagnostics								
CIPS test	−3.5010**	−3.3390*	−4.4400***	−4.6890***	−2.6510	−2.6480	−4.0500**	−3.5600**
ΔCIPS	−5.2160***	−5.0260***	−5.7320***	−5.9160***	−4.4580***	−4.1270**	−5.4080***	−4.8660***
CD tests	21.3190***	14.5140***	1.6060	1.6390	11.6340***	14.8190***	1.3410	1.6260
Panel C: 1990–2007								
trade	0.1983**	0.2261**	0.2398*	0.3218**	0.0369	−0.0302	0.0749	0.0320
	(0.0929)	(0.0913)	(0.1310)	(0.1385)	(0.0578)	(0.0500)	(0.0495)	(0.0212)
fopen	−0.1778	−0.1477	0.1229**	0.1344***	−0.0362	−0.0036	−0.0644**	−0.0372*
	(0.1091)	(0.1149)	(0.0521)	(0.0371)	(0.0233)	(0.0406)	(0.0285)	(0.0209)
Diagnostics								
CIPS test	−4.7520***	−4.3660***	−5.8860***	−5.1650***	−4.2350**	−3.6010**	−4.4770**	−4.6220***
ΔCIPS	−3.3590*	−3.4550*	−4.3280**	−3.4280**	−3.4480**	−3.2220*	−4.1060*	−4.0160**
CD tests	17.0860***	26.0650***	1.2390	1.6300	17.4620***	22.4290***	1.4920	1.3090

Note: The values in parentheses are the standard errors based on a Newey–West type variance estimator.
***,**,* Significant at the 1%, 5% and 10% level, respectively.

whereas those of the trade variables are not statistically significant for the GDP-deflator-based inflation regression but significantly positive at the conventional level for the CPI-based inflation regression. However, from the diagnostic tests, MG yields stationary but cross-section-dependent residuals, indicating that the MG estimation is likely to produce misleading inferences.

We then carry out the CCEMG estimates that control for cross-section dependence. As can be seen, the diagnostic tests are sound in that residuals are stationary and cross-section-independent, clearly supporting the use of CCEMG estimators for the developing country data. These preferred models suggest that the average impact of openness differs across countries and show that the estimates for trade openness are positive and highly significant, implying that greater international trade strengthens inflation, in general. Similarly, financial openness, on the average, reinforces inflation, as the estimates on financial openness are also positive and statistically significant. The effect of globalization is economically meaningful as well. Increasing trade (financial) openness by one percent increases average inflation by 0.11 to 0.20 (0.09 to 0.12) percent, depending upon alternative inflation measures. Overall, the data indicate that openness does not seem to be a constraint on policy makers' incentives to inflate. Instead, openness intensifies the incentive of policy makers to inflate. As a result, monetary policy tends to be more inflationary in more open developing economies. This is in sharp contrast to theoretical predictions that emphasize the role of globalization, either trade or financial openness, as a discipline device for central banks and hence that globalization is associated with lower inflation [e.g., Romer (1993); Lane (1997); Gruben and McLeod (2002)].[5] A possible explanation is that financial globalization exposes firms to large external shocks that may undermine the credibility of monetary authorities or lead to sudden reversals of capital inflows, leading to devaluations and higher inflation. As put forth by Rodrik (1998), capital account liberalization makes developing economies that do not have the right infrastructures more vulnerable to destabilizing, inflationary capital flows. Also, recent crises experienced by developing countries suggest that the inherently volatile capital flows, which are manifest most severely in sudden stops [Calvo and Reinhart (2002)], hot money [Stiglitz (1999)], and even capital flight [Aghion et al. (2004)], lead to severe recessionary consequences, especially during economic downturns in countries with low absorptive capacity and weak institutions [Aghion et al. (1999)]. Similarly, trade liberalization that leads to higher integration of world goods markets may make developing countries in particular more vulnerable to crises because of their production specialization, nondiversified sources of income, unstable policies, and/or weak institutions [e.g., Loayza and Raddatz (2007)]. Hence, higher trade and financial flows may actually reinforce, rather than reduce, inflation in the process of globalization of developing countries.

Regarding the developed country data, the MG results fail to pass the cross-section-independence test. However, diagnostics are sound in the case of the CCEMG results, as expected. Residuals are stationary and cross-section-independent, suggesting preference for the CCEMG estimator over MG. The CCEMG estimates for both trade and financial openness are not statistically significant, indicating that, on the average, globalization does not exert a significant effect on inflation in developed economies. This finding calls into question the assertion that globalization has significantly affected the inflation performance of highly developed countries [Lane (1997); Gruben and McLeod (2004)] but accords with the findings of Romer (1993) and Badinger (2009). The most likely explanation is that the time-inconsistency problem has been successfully resolved by this group of countries through the creation of a proper institutional framework for central banks, as suggested by Romer (1993) and Badinger (2009).

3.2. Model Stability

We now turn to the analysis of subsample stability with respect to the time dimension. We spilt the time span (1970–2007) into two subperiods: 1970–1989 and 1990–2007. This split is critical. On one hand, the empirical findings of most

https://doi.org/10.1017/S1365100514000510 Published online by Cambridge University Press

previous cross-country investigations [e.g., Romer (1993); Lane (1997); Terra (1998)] that openness leads to lower long-term inflation rates have focused on the experience during the 1970s and 1980s when monetary policy is more characterized by discretion or a lack of commitment and when inflation is more volatile. On the other hand, it is argued that the inverse link of inflation with openness has strengthened since the 1990s [Gruben and McLeod (2004)] in which the economies are more integrated [Ball (2006); Lane and Milesi-Ferretti (2007)], inflation is less volatile [International Monetary Fund (2006)], and monetary policy is more characterized by rules, instead of discretion [Evans (2012)]. The subperiod 1990–2007, includes the globalization era, covers the Great Moderation, and captures the adoption of inflation targeting, implicitly or explicitly. Therefore, such an experiment with subperiods not only serves as a comparison but also avoids the issue of potential structural change in inflation variability.

Panels B and C of Table 2 report the respective results for the period 1970–1989 and 1990–2007. Focusing on the first subperiod in Panel B, the MG estimator produces either nonstationary or cross-section-dependent residuals, whereas the residuals from the CCEMG estimator are stationary and cross-section-independent in both the developing and developed country samples.

Based on our preferred CCEMG models, we find on the average that neither trade nor financial openness impose any restriction on inflationary policy in the sample of developing countries, as the respective estimates on trade and financial openness variables are of no statistical significance. In the developed country sample, greater trade does constrain inflation, as predicted by Lane (1997). The estimate on trade is negative and statistically significant at the 5% level. However, financial openness has no significant impact on inflation, on the average. Because most developed countries and some emerging market economies start to liberalize their financial accounts on a large scale from the early 1990s onward, it is not surprising to find a nonsignificant effect of financial openness for both country groups during this period of time.

Regarding the period 1990–2007 in Panel C, although the MG estimator produces nonstationary but cross-section-dependent residuals, the residuals from the CCEMG estimator are stationary and cross-section-independent in both the developing and developed country samples. This validates the use of the CCEMG estimator.

The CCEMG estimates for trade and financial openness variables are positive and statistically significant for the developing country sample. This implies that inflation will rise if a country opens up its trade and/or financial account. For the case of developed countries, the CCEMG estimates for trade openness are not statistically significant. It is consistent with Bleaney (1999) that the inverse relationship between trade openness and inflation disappears from the early 1990s on, but only for the developed country case. On the other hand, the financial openness variable is negative and statistically significant. Financial openness, in general and on the average, constrains inflation in the developed countries. The evidence is in line with the finding of Obstfeld (1998) that increased international capital mobility could have a disciplining effect on monetary policy, and of Gruben and McLeod (2002) that increased international asset substitutability reduces the effectiveness of using inflation as a source of government revenues, thereby reducing the inflationary pressure on central banks. However, our data indicate that this is relevant only for developed countries.

3.3. Individual Long-Run Estimates

Another important point is that the positive long-run effect of openness on inflation for the developing country group during the period 1970–2007 is the average of the effects in individual countries and may therefore mask important differences across countries. As mentioned, the effects of openness on inflation may be highly heterogeneous, because of cross-country differences in income levels, policies, and various other characteristics. Similar logic also applies for the case of developed countries. To examine heterogeneity in the long-run effects of openness on inflation across countries, we present the individual country CCEMG estimates of the coefficients on trade and financial openness in Table 3. These estimates should be interpreted with caution given the relatively short time dimension of our data. Nevertheless, it can be concluded safely that there is considerable heterogeneity in the long-run effects of openness on inflation across countries.

For the developing countries in Panel A, the trade coefficients range from -2.116 (Zimbabwe) to 3.549 (Russia). The coefficients of financial openness lie in the range between −1.781 (Brazil) and 3.025 (Bolivia). Panel A indicates that, although the long-run effect of openness on inflation is positive in general or on the average for developing countries, openness does not have a positive long-run effect on inflation in all developing countries. More specifically, we find that an increase in trade (financial) openness is associated with an increase in inflation in 61 (52) out of 83 developing countries, with 33 (29) of the positive coefficients being statistically significant. On the other hand, an increase in trade (financial) openness is associated with a decrease in inflation in 22 (31) developing countries; only 7 (13) of the negative coefficients are statistically significant.

In the case of developed countries, Panel B of Table 3 shows that the trade coefficients range from −1.250 (Estonia) to 0.334 (Bahrain), and the coefficients of financial openness falls in the range between −0.262 (United Kingdom) and 0.421 (Poland). This finding indicates that, although the long-run effect of openness on inflation is weak in general or on average for developed countries, openness does have a positive or negative long-run effect on inflation in some of them. In particular, we find that an increase in trade (financial) openness is associated with an increase in inflation in 25 (19) out of 39 developed countries, with 11 (5) of the positive coefficients being statistically significant. On the other hand, an increase in trade (financial) openness is associated with a decrease in inflation in 14 (20) developed countries, with 6 (8) of the negative coefficients being statistically significant.

TABLE 3. CCEMG individual estimates of trade and financial openness on inflation

	trade	*fopen*		*trade*	*fopen*		*trade*	*fopen*		*trade*	*fopen*
Panel A: Developing countries (83)											
Algeria	0.1810	0.0870	Dominica	0.0530	−0.0170	Latvia	2.8230***	−0.7410	Philippines	0.1280**	0.3200***
Argentina	−0.5590	1.7360***	Dominican Rep.	0.1790**	0.3180***	Lesotho	−0.0560	0.0220	Romania	−0.9800	1.4820***
Bangladesh	−0.2930	−0.3660**	Ecuador	0.0830	−0.0360	Lithuania	−0.8330***	1.9800**	Russian Federation	2.6610**	−0.1400
Belarus	0.4610	0.4720***	Egypt	0.0720	−0.0040	Macedonia, FYR	1.0740**	1.2230**	Rwanda	−0.2160**	0.0940
Benin	−0.0400	0.0710***	El Salvador	0.0340	−0.0350***	Madagascar	−0.0620	0.1740***	Senegal	0.0720	0.1600***
Bhutan	0.0730	−0.0060	Ethiopia	−0.1810***	−0.0710*	Malawi	−0.0520	0.1110**	South Africa	0.2090***	−0.0330
Bolivia	1.1290	3.1850***	Fiji	0.0840	−0.1820***	Malaysia	0.0340	−0.0980***	Sri Lanka	0.0100	0.0360
Botswana	−0.0340	0.0620	Gabon	0.0930	−0.1160	Mali	−0.0250	0.1020*	Swaziland	−0.1560***	0.0510
Brazil	0.3840	−0.5000	Ghana	−0.0130	−0.1420	Mauritania	−0.0430	−0.0630	Tanzania	0.0030	0.0290
Bulgaria	−0.8470	0.0370	Grenada	−0.1100*	−0.2930***	Mauritius	0.1170*	0.0740**	Thailand	0.0730	−0.0280
Burundi	−0.1690***	0.0550	Guatemala	−0.0300	0.1730***	Mexico	−0.1810*	0.7630***	Togo	−0.2190**	0.1260***
Cameroon	0.1130***	0.0330***	Haiti	−0.0540	0.0160	Morocco	−0.0690	0.0700*	Tunisia	0.1490**	0.0400
Cape Verde	−0.3820	0.0600	Honduras	0.1210***	0.2060***	Myanmar	−0.0770***	−0.0210	Turkey	−0.1280	−0.1320
Central African Rep.	0.1220	0.0090	India	−0.0450	0.0150	Nepal	−0.0090	−0.0920	Uganda	0.8110***	−0.6060***

TABLE 3. Continued.

	trade	*fopen*		*trade*	*fopen*		*trade*	*fopen*		*trade*	*fopen*
Chad	−0.0020	0.0600	Indonesia	0.4360***	0.1640***	Niger	0.0200	0.0270	Ukraine	0.4600	−0.0490
Chile	−0.1330	0.6390	Iran, Rep.	0.0810**	−0.0040	Nigeria	0.1750*	0.0750	Uruguay	−0.1360	0.2130
China	0.1230***	−0.0460*	Jamaica	0.2700	−0.0800	Pakistan	−0.0750	0.0730***	Venezuela, RB	0.2870	0.3080***
Colombia	0.0100	−0.1260***	Jordan	0.1470	0.0570	Panama	0.0800*	−0.0190	Yemen, Rep.	0.1030	0.2170*
Congo, Rep.	0.0530	0.0510	Kazakhstan	−1.2550***	−0.0820	Papua New Guinea	−0.3150*	0.0850	Zambia	0.6630***	−0.0670
Costa Rica	0.3960***	0.2640***	Kenya	−0.1100	0.2910***	Paraguay	−0.0480	0.1450***	Zimbabwe	−0.1640	0.1230***
Cote d'Ivoire	0.2930**	−0.1100**	Lao PDR	0.5750	−0.3410	Peru	1.7790**	−1.2930**			
Panel B: Advanced countries (39)											
Australia	−0.1060	−0.0380	France	0.0920	−0.0450***	Korea, Rep.	0.0200	0.0100	Singapore	0.0510	−0.0160
Austria	0.0990***	0.0000	Germany	0.0180	−0.0110	Kuwait	−0.5550	−0.0920***	Slovak Rep.	0.0900	0.1840
Bahrain	0.1710	0.0130	Greece	0.3200***	0.0180	Luxembourg	−0.3050***	−0.0930	Slovenia	1.0780**	−0.8070**
Belgium	0.1350*	−0.0430***	Hong Kong	0.1280***	−0.1560**	Malta	0.0930**	0.0010	Spain	−0.2080***	0.0240
Canada	−0.0690***	0.0690**	Hungary	−0.3330**	0.0360	Netherlands	−0.0890	0.1050***	Sweden	−0.0440	−0.0600***
Cyprus	−0.1060***	0.0050	Iceland	−0.1570	0.0630**	New Zealand	−0.0730	−0.0710	Switzerland	0.1140**	−0.0180
Czech Rep.	0.2930	−0.3820*	Ireland	−0.0750	0.0540***	Norway	0.1770	0.0570**	Trinidad and Tobago	0.1450***	0.0020
Denmark	0.1630***	−0.0430	Israel	0.1160	0.2900	Poland	0.0750	0.3080***	United Kingdom	0.1600**	−0.1600**
Estonia	−0.6790***	−0.0190	Italy	0.1370***	−0.0470***	Portugal	0.0350	−0.0120	United States	−0.0180	−0.0240
Finland	−0.0570	0.0330	Japan	0.0550*	−0.0280*	Saudi Arabia	−0.2210	−0.0610			

***,**,* Significant at the 1%, 5% and 10% levels, respectively.

TABLE 4. CCEMG estimation of error correction models

	Developing countries		Advanced countries	
	inf_gdp	*inf_cpi*	*inf_gdp*	*inf_cpi*
		Panel A: 1970–2007		
ϕ	−0.2771***	−0.5322***	−0.7487***	−0.6711***
	(0.0748)	(0.1575)	(0.0908)	(0.0815)
$\triangle$*trade*	0.0429	0.0543	−0.0224	0.0355
	(0.0784)	(0.1102)	(0.0360)	(0.0326)
$\triangle$*fopen*	0.0071	0.1302	0.0251	−0.0001
	(0.0726)	(0.1066)	(0.0270)	(0.0165)
		Panel B: 1970–1989		
ϕ	−0.4124**	−0.4541***	−0.2860**	−0.4813***
	(0.1677)	(0.1644)	(0.1152)	(0.1837)
$\triangle$*trade*	−0.1459	−0.1816	0.0428	−0.0913
	(0.1852)	(0.1949)	(0.0744)	(0.0971)
$\triangle$*fopen*	0.0976	0.0782	−0.1087	−0.0033
	(0.1184)	(0.1231)	(0.0884)	(0.0847)
		Panel C: 1990–2007		
ϕ	−0.4925***	−0.8628***	−0.3819**	−0.9175***
	(0.1037)	(0.1421)	(0.1573)	(0.2031)
$\triangle$*trade*	0.1168	0.0334	0.0722	0.0934
	(0.1069)	(0.0883)	(0.0890)	(0.0657)
$\triangle$*fopen*	−0.0458	0.1475**	0.0267	−0.0495*
	(0.1362)	(0.0698)	(0.0459)	(0.0290)

Note: The values in the parentheses are the robust standard errors.
***,**,* Significant at the 1%, 5% and 10% levels, respectively.

3.4. Short-Run Dynamics

Given the very nature of the time-inconsistency problem, Alfaro (2005) claims that a more relevant exercise is to assess whether openness serves as a commitment mechanism for restricting inflation in the short run. She finds that the estimate for trade openness is, in the short run, not significant for high-income countries but significantly positive for middle- and low-income countries, concluding that openness does not play a role in restricting inflation in the short run. As a final check, therefore, we consider whether both opennesses serve as commitment mechanisms for restraining inflation in the short run.

Table 4 reports the error correction models attached to the CCEMG estimate, which has shown a cointegration relation between the variables. The coefficient attached to $y_{it-1} - \beta_i' x_{it-1}$(i.e., the error correction term ϕ) measures the speed of adjustment of inflation to a deviation from the long-run equilibrium relation between inflation and its determinants. As expected, across alternative inflation indicators, different model specifications, and various time periods, the coefficient of the error correction term ϕ is negative and significant, from which it can be

https://doi.org/10.1017/S1365100514000510 Published online by Cambridge University Press

concluded that the variables are cointegrated. In the short run, however, trade openness tends to have no significant effect on inflation for either the developing or the developed country samples. The coefficient estimates of $\Delta trade$ are not statistically significant. As for financial openness, the estimates are also not statistically significant, except for the CPI-based inflation regression models in the period 1990–2007, where changes in financial openness are significantly positive in the sample of developing countries but significantly negative in the developed country sample. Overall, the data imply that both forms of openness tend not to restrict inflation. The result for the whole sample does, however, not imply that both types of openness do not discipline monetary policy in each individual country.

4. EXPLAINING DIFFERENTIAL LONG-RUN IMPACTS OF OPENNESS ON INFLATION

The cross-country differences in the long-run effect of openness on inflation pose a new question: what factors can explain this heterogeneity or, in other words, what factors determine the long-run impact of openness on inflation? A potential way to answer this question is to explore whether the observed pattern of long-run effects of openness can be linked to cross-country differences in the quality of political institutions, the extent of central bank independence, the exchange-rate regimes, the degree of financial development, the real interest rates, and the legal system. These variables have been proven among determinants of inflation, trade, or financial openness. We thus regress the trade and financial openness coefficient estimates from the CCEMG model on these variables.

The quality of political institutions is measured by the extent of democratization using the Polity IV data set [Marshall and Jaggers (2010)]. Polity IV measures a political regime by the polity score, which ranges from –10 (strongly autocratic) to +10 (strongly democratic). This measure reflects the degree of competitiveness in political participation, the openness and competitiveness in the selection of the legislature, and the constitutional constraints on executive powers. It also incorporates subjective information on checks and balances on executive powers, the degree of restrictions on electoral participation, and the extent to which political participation is regulated. The related literature has found that democracy is associated with lower inflation rates [Desai et al. (2003)], higher trade flows [Milner and Kubota (2005); Yu (2010)], and greater cross-border financial flows [Aizenman (2004); Campos and Coricelli (2012)]. Romer (1993) even finds that the inverse link between inflation and trade openness is stronger in countries that are politically unstable.

Central bank independence is an effective means of ensuring price stability [Meade and Crowe (2007); Cukierman (2008)]. An independent central bank can place full priority on low levels of inflation, whereas in countries with a more dependent central bank, other considerations (notably, reelection perspectives of politicians and a low level of unemployment) may interfere with the objective of price stability. Romer (1993) finds that central bank independence weakens the negative link between inflation and trade openness. Central bank independence is proxied by the actual turnover rate of central bank governors obtained from Dreher et al. (2011).[6] A high turnover rate indicates fragile sovereignty of authority and minimizes the opportunity of the central bank to adopt strategies.

Adopting fixed-exchange-rate regimes creates incentives for policy makers to control money supply and restricts inflation [Alfaro (2005)]. Exchange-rate regimes with lower uncertainty and transaction costs—namely, conventional pegs and currency unions—are significantly more pro-trade than flexible regimes [Egger (2002); Klein and Shambaugh (2006); Adam and Cobham (2007)]. Further, Hoffmann and Tillmann (2012) show that financial integration increases the national price level under managed-exchange-rate regimes and lowers the price level under floating exchange rates. Bowdler (2009) finds that the negative response of output to inflation increases with exchange rate flexibility. We use a new exchange-rate-regime classification constructed by Rogoff and Reinhart (2004), which considers parallel exchange-rate data when assessing whether a country has, de facto, maintained a pegged or a flexible regime.[7] A higher value reflects a more floating exchange rate system.

A more developed financial sector may increase the scope of action of policy, resulting in improved policy performance. Posen (1995), for example, argues that the central bank can guarantee price stability only as long as the financial sector is ready to support policies associated with reducing inflation. More developed financial sectors will lead to more successful stabilization policies. Better macroeconomic conditions further support the development of financial systems. Whereas Cecchetti and Krause (2001) find evidence that an improvement in the depth of the financial sector and the intermediation process leads to reduction in inflation and output variability, Kim and Lin (2010) show that lower inflation is beneficial for financial sector development. Moreover, financial development as a source of comparative advantage and an insurance mechanism determines the pattern and flows of international trade [Svaleryd and Vlachos (2002); Beck (2003); Kim et al. (2010)]. Local financial market development also determines whether a country can benefit from financial openness [Hermes and Lensink (2003); Alfaro et al. (2004)]. Financial development is proxied by private credit, the claims on private sector by deposit money banks and other financial intermediaries as a share of GDP from Beck et al. (2000), and is probably the most important banking development measure because it reflects the extent to which any entrepreneur or company with a sound project can obtain bank finance.

The real interest rate is obtained from the World Development Indicators of the World Bank (2012). Interest rates drive inflation dynamics, not only through demand, but also through supply. Chowdhury et al. (2006), Ravenna and Walsh (2006), and Tillmann (2008) show, for instance, that higher interest rates translate into higher marginal costs of production and, eventually, into higher inflation. Blankenau et al. (2001) also finds that world real interest rate shocks are responsible for more than half of the fluctuations in net exports and net foreign assets.

Legal systems differ systematically in their enforcement of contracts [Djankov et al. (2003); La Porta et al. (2008)]. Common law systems are better at enforcing contracts than civil law systems. This is due, in part, to the fact that civil law states heavily regulate legal proceedings. Because legal systems differ in their enforcement of contracts, domestic legal institutions could influence patterns of international trade] Helpman (2006) Nunn (2007) find that countries with legal systems that are systematically better at contract enforcement will enjoy higher levels of trade. The legal origin indicator takes on the value of one for common law countries and zero for non-common law countries. The data on legal origins are taken from La Porta et al. (1999). These variables are averaged over the period from 1970 to 2007. The exception is the data on legal origins.

Table 5 reports the estimates. The OLS estimates in Model (1) indicate that the long-run effect of trade on inflation is significantly positively associated with the exchange-rate regimes and real interest rates. In contrast, there is no statistically significant correlation of the long-run trade effect with political institutions, central bank independence, financial development, or legal origins. On the other hand, the OLS estimates in Model (6) show that the long-run effect of financial openness on inflation is significantly positively linked to central bank independence but negatively related to legal origins. There is no statistically significant association of the long-run effect of financial openness with the quality of political institutions, the exchange-rate regimes, the level of financial development, or real interest rates.

In the presence of endogeneity, however, OLS estimation would produce incorrect inference. The higher sensitivity of inflation to openness might increase the pressure and the incentive to reduce monetary discretion (by committing to a fixed-exchange-rate regime or other monetary policy rules), increase the level of central bank independence, and/or implement structural reforms (such as financial or political reforms) that increase the cost of monetary discretion and therefore reduce the central bank's incentive to exploit output gains from unexpected inflation. That said, there might be feedback effects from the quality of political institutions, central bank independence, the exchange-rate regimes, financial development, and real interest rates.[8]

We then run two-stage least squares (TSLS) regressions that explicitly control for simultaneity bias and reverse causality. In particular, we use endowments (a dummy for oil exporters), geographic factors (such as latitude, landlock, and tropics), and cultural features (such as religions and ethnic fractionalization) as instruments.[9] These endowments, geographics, and culture shape a country's institutional development, such as financial development [La Porta et al. (1999); Acemoglu et al. (2001); McCaig and Stengos (2005)], the extent of democratization [Barro (1999); Acemoglu et al. (2001)], the level of central bank independence [Acemoglu et al. (2008)], and exchange-rate regimes [Bernhard and Leblang (1999); Aghion et al. (2009)]. Instrumental variables satisfy two requirements: they must be correlated with the endogenous regressors, and they must be orthogonal to the error process. Therefore, to check for the quality of these instruments, we perform the Sargan test and the Hansen J-test, which is consistent in the

TABLE 5. Long-run inflation effects and country-specific factors

	Trade coefficient estimates					Financial openness coefficient estimates				
	(1) OLS	(2) TSLS	(3) TSLS	(4) TSLS	(5) CUE	(6) OLS	(7) TSLS	(8) TSLS	(9) TSLS	(10) CUE
Political institutions	−0.0027	−0.0007	−0.0015	−0.0038	0.0051	−0.0075	−0.0161**	−0.0225*	−0.0204*	−0.0200*
	(0.0198)	(0.0126)	(0.0077)	(0.0087)	(0.0120)	(0.0134)	(0.0089)	(0.0122)	(0.0111)	(0.0115)
Central bank independence	0.4872	0.9006**	1.0101***	0.9861**	0.1443**	2.0988***	3.2253***	3.1111***	3.3014***	2.3438***
	(1.0215)	(0.4326)	(0.3911)	(0.3963)	(0.0713	(0.9587)	(1.0571)	(0.9016)	(1.0570)	(0.5731)
Exchange-rate regimes	0.3506***	0.1339**	0.1446*	0.1605**	0.9899*	0.0002	−0.3280**	−0.2735**	−0.3426**	−0.2077*
	(0.1228)	(0.0684)	(0.0877)	(0.0753)	(0.5640)	(0.1372)	(0.1676)	(0.1342)	(0.1655)	(0.1105)
Private credit	−0.0089	−0.0389	−0.0136	−0.0399	−0.0446	0.0346	0.2152**	0.2320**	0.2439**	0.1524*
	(0.1399)	(0.0811)	(0.0668)	(0.0705)	(0.0774)	(0.1188)	(0.1057)	(0.1053)	(0.1158)	(0.0814)
Real interest rates	0.0810**	0.0465	0.0296	0.0340	0.0286	−0.0003	0.0388	0.0211	0.0375	0.0190
	(0.0394)	(0.0382)	(0.0309)	(0.0307)	(0.0331)	(0.0064)	(0.0302)	(0.0183)	(0.0270)	(0.0211)
Legal origin dummy	0.0489	0.1759	0.0756	0.0873	0.0865	−0.2589**	−0.1864*	−0.1909*	−0.2420*	−0.2111*
	(0.2124)	(0.1331)	(0.1189)	(0.1179)	(0.1295)	(0.1206)	(0.1025)	(0.1144)	(0.1445)	(0.1213)
Constant	−0.8980	−0.3481	−0.3962	−0.3594	−0.3114	−0.2351	−0.5885	−0.6378*	−0.6453	−0.3833
	(0.6648)	(0.4192)	(0.3745)	(0.3498)	(0.3947)	(0.4837)	(0.3811)	(0.3672)	(0.3978)	(0.3086)
R^2	0.2535	0.3755	0.3330	0.3514	0.3898	0.3248	0.3195	0.4792	0.3539	0.4881
Hansen J-statistic		0.6894	0.0974	0.5514	2.683		1.4849	0.424	1.338	0.373
		[0.7084]	[0.7549]	[0.7590]	[0.1014]		[0.6857]	[0.5149]	[0.7201]	[0.5412]
Sargan test statistic		0.473	0.101	0.598			2.583	2.491	2.807	
		[0.7893]	[0.7502]	[0.7414]			[0.4604]	[0.4769]	[0.4223]	
Durbin–Wu–Hausman test		8.7533	4.4203	4.6136	4.8318		2.9555	3.1193	3.2983	3.8432
		[0.0125]	[0.0961]	[0.0995]	[0.0892]		[0.0855]	[0.0773]	[0.0693]	[0.0499]

Note: The values in parentheses (brackets) are the robust standard errors (p-values) of corresponding coefficient estimates.
***,**,* Significant at the 1%, 5%, and 10% levels, respectively.

presence of the general form of heteroskedasticity, to validate that the instruments are orthogonal to the error process.

TSLS estimates in Model (2) indicate that the long-run effect of trade on inflation is significantly positively associated with central bank independence and the exchange-rate regimes. The estimate on real interest rates is positive but not statistically significant. Those on political institutions, financial development, and legal origins remain nonsignificant. This finding indicates that the impact of trade on inflation is more negative for countries with higher degrees of central bank independence and/or less flexible exchange-rate regimes. On the other hand, as illustrated in Model (7), also from the TSLS estimator, the long-run effect of financial openness on inflation is significantly positively associated with central bank independence and financial development, but negatively related to the quality of political institutions, the exchange-rate regimes, and legal origins. The real interest rate variable turns positive but remains nonsignificant. The influence of financial openness is more negative for countries with higher quality of political institutions, more independent central banks, more flexible exchange rate systems, less developed financial systems, and/or common law.

Indeed, a potential problem with this analysis is that the estimated coefficients on trade (and financial openness) are not statistically significant for all countries. However, the results do not change qualitatively if we set the long-run effect of openness on inflation equal to zero for the countries with insignificant coefficients [models (3) and (8)] and exclude countries with insignificant coefficients [models (4) and (9)]. Finally, we confirm these results with the continuously updated GMM estimator (CUE) in models (5) and (10).

With respect to instruments, both the Sagan statistic and the Hansen J-statistic are far from a rejection of its null hypothesis that the full set of orthogonality conditions is valid. The evidence thus indicates that the instruments used are appropriate. Moreover, the Durbin–Wu–Hausman test rejects the null of exogeneity. Overall, the data seem to suggest that cross-country differences in the long-run effect of openness on inflation can be at least partly explained by cross-country differences in the quality of political institutions, central bank independence, the exchange rate system, financial development, and/or legal traditions.

5. CONCLUDING REMARKS

This paper examines whether there exists a long-run relationship between globalization and inflation. In particular, we assess the nonstationarity and cointegration properties between openness and inflation. This is done in a panel data context controlling for both cross-section dependence and heterogeneity in a common-factor model. Some important results emerge. Both trade and financial globalization exert a significant and positive effect on inflation in a sample of developing countries. The data suggest that globalization may not impose a disciplinary impact on the central banks of developing countries. Instead, globalization may strengthen the fear of heightening macroeconomic instability of these countries in the process of liberalizing either trade or financial accounts, as the liberalization generally comes at the cost of higher inflation. As for advanced countries, there is, on the average, no significant impact of openness on inflation.

However, the inverse link between trade or financial globalization and inflation cannot be ruled out under all circumstances. In particular, our cross-country regression results show that the long-run impact of trade on inflation is more negative for countries with greater degrees of central bank independence and/or less flexible exchange-rate regimes, and that the long-run influence of financial openness is more negative for countries with a higher quality of political institutions, more independent central banks, more flexible exchange-rate regimes, less developed financial systems, and/or British common law.

NOTES

1. For example, whereas Gruben and McLeod (2002) estimate the effect of capital account restrictions on inflation, Badinger (2009) and Spiegel (2009) consider de facto financial openness.
2. Data on inflation and trade are taken from the World Development Indicators of the World Bank (2012).
3. They construct estimates of external assets and liabilities for 145 countries over the period 1970 to 2007. This is why our sample period ends in 2007.
4. It is noted that, from Table B.1, the pooled estimators are suggested to be severely biased, given the diagnostic tests indicating that residuals are cross-section-dependent and/or nonstationary. This bias may arise from the misspecification of homogeneity, nonstationarity, and/or the failure to account for unobserved common factors appropriately. We therefore emphasize the estimates from heterogeneous models.
5. However, the evidence that trade increases a country's long-run incentive to create inflation is consistent with Cavelaars (2009) and Cooke (2010), who emphasize the international expenditure-switching effect, or Evans (2012), who operates through the beggar-thy-neighbor channel.
6. The reason for the use of the turnover rate is that legal measures of central bank independence may not reflect the true relationship between the central bank and the government. Especially in countries where the rule of law is less strongly embedded in the political culture, there can be wide gaps between the formal, legal institutional arrangements and their practical impact. This is particularly likely to be the case in many developing economies.
7. Their 14 categories of exchange-rate regimes are (1) no separate legal tender, (2) preannounced peg or currency board arrangement, (3) preannounced horizontal band that is narrower than or equal to $\pm 2\%$, (4) de facto peg, (5) preannounced crawling peg, (6) preannounced crawling band that is narrower than or equal to $\pm 2\%$, (7) de facto crawling peg, (8) de facto crawling band that is narrower than or equal to $\pm 2\%$, (9) preannounced crawling band that is wider than or equal to $\pm 2\%$, (10) de facto crawling band that is narrower than or equal to $\pm 5\%$, (11) moving band that is narrower than or equal to $\pm 2\%$ (i.e., allows for both appreciation and depreciation over time), (12) managed floating, (13) free floating, and (14) free falling.
8. We thank the anonymous referee for pointing this out. We hence treat all these variables as potentially endogenous, except for the legal origin variable.
9. These instrumental variables are obtained from the Global Development Network Growth Database and La Porta et al. (1999).

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APPENDIX A: PANEL UNIT ROOT AND CROSS-SECTION-DEPENDENCE TESTS

A.1. TESTING FOR PANEL UNIT ROOTS

Consider the pth-order augmented Dickey–Fuller (ADF) regression:

$$\Delta q_{it} = a_i + b_i q_{it-1} + c_i t + \sum_{j=1}^{p} d_{ij} \Delta q_{it-j} + \varepsilon_{it}, \qquad \textbf{(A.1)}$$

where q is the logarithm of inflation, the logarithm of the jth regressor $x_{j,it}$, or regression residuals from equation (1). ε_{it} are errors. In testing for unit roots, the null hypothesis is

$$H_o : b_i = 0, \quad i = 1, \ldots, N \qquad \textbf{(A.2)}$$

against the alternative that

$$H_o : b_i < 0, \quad i = 1, \ldots, N_1; \quad b_i = 0, \quad i = N_1 + 1, \ldots, N, \qquad \textbf{(A.3)}$$

where N_1 is such that N_1/N is nonzero and tends to be a constant as N goes to infinity. Pesaran (2007) proposes to test (A.2) against (A.3) by computing the simple average of the t-ratios of the OLS estimates of b_i in the Dickey Fuller regression augmented with the cross-section averages $\bar{q}_{t-1}$ and $\Delta\bar{q}_{t-j}$,for $j = 0, \ldots, p$:

$$\Delta q_{it} = a_i + b_i q_{it-1} + c_i t + \sum_{j=1}^{P} d_{ij} \Delta q_{it-j} + g_i \bar{z}_t + e_{it}, \qquad \textbf{(A.4)}$$

where $\bar{z}_t = (\bar{q}_{t-1}, \Delta\bar{q}_t, \Delta\bar{q}_{t-1}, \ldots, \Delta\bar{q}_{t-p})'$, as

$$\text{CIPS} = \frac{1}{N}\sum_{i=1}^{N}\tilde{t}_i, (\boldsymbol{A}) \tag{A.5}$$

where $\tilde{t}_i$ is the OLS t-ratio of b_i.

The critical values for the CIPS tests are given in Tables 2(a)–2(c) in Pesaran (2007).

A.2. CROSS-SECTION-DEPENDENCE TESTS

To check for cross-section dependency, we apply a cross-section-dependency (CD) test proposed by Pesaran (2004). Pesaran (2004) showed that his CD test can also be applied to a wide variety of models with small/large N and T. Additionally, this simple diagnostic test does not require an a priori specification of a connection or spatial matrix. The CD test is based on a simple average of all pairwise correlation coefficients of the OLS residuals from the individual regressions in the panel,

$$\rho_{ij} = \rho_{ji} = \frac{\sum_{t=1}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{jt}}{\left(\sum_{t=1}^{T}\hat{\varepsilon}_{it}^2\right)^{1/2}\left(\sum_{t=1}^{T}\hat{\varepsilon}_{jt}^2\right)^{1/2}},$$

where the $\hat{\varepsilon}_{it}$ is the OLS estimate of ε_{it} from the individual-country regressions. The proposed CD test by Pesaran (2004) is then given by

$$CD = \sqrt{\frac{2T}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\rho_{ij}. \tag{A.6}$$

The CD test statistic has exactly zero mean for fixed values of T and N, under a broad class of panel data models.

For unbalanced data,

$$CD = \sqrt{\frac{2}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\sum\sqrt{T_{ij}}\rho_{ij} \tag{A.7}$$

where $T_{ij} = \#(T_i \cap T_j)$, the number of common time-series observations between units i and j,

$$\rho_{ij} = \frac{\sum_{t\in T_i\cap T_j}(\hat{\varepsilon}_{it} - \bar{\hat{\varepsilon}}_i)(\hat{\varepsilon}_{jt} - \bar{\hat{\varepsilon}}_j)}{\{\sum_{t\in T_i\cap T_j}(\hat{\varepsilon}_{it} - \bar{\hat{\varepsilon}}_i)^2\}^{1/2}\{\sum_{t\in T_i\cap T_j}(\hat{\varepsilon}_{jt} - \bar{\hat{\varepsilon}}_j)^2\}^{1/2}},$$

and

$$\bar{\hat{\varepsilon}}_i = \frac{\sum_{t\in T_i\cap T_j}\hat{\varepsilon}_{it}}{\#(T_i \cap T_j)}.$$

The modified statistic accounts for the fact that the residuals for subsets of t are not necessarily zero-mean.

https://doi.org/10.1017/S1365100514000510 Published online by Cambridge University Press

APPENDIX B: POOLED ESTIMATES OF THE LONG-RUN EFFECT OF OPENNESS ON INFLATION

TABLE B.1. Pooled estimates of the long-run effect of openness on inflation

	Developing countries				Advanced countries			
	FE		CCEP		FE		CCEP	
	inf_gdp	*inf_cpi*	*inf_gdp*	*inf_cpi*	*inf_gdp*	*inf_cpi*	*inf_gdp*	*inf_cpi*
	Panel A: 1970–2007							
trade	−0.0575***	−0.0450**	−0.0214	0.0089	0.0618**	0.0296	0.0310	0.0588**
	(0.0191)	(0.0201)	(0.0296)	(0.0331)	(0.0303)	(0.0330)	(0.0358)	(0.0270)
fopen	−0.0166	−0.0177	0.0964***	0.1029***	−0.0549***	−0.0432***	−0.0023	0.0094
	(0.0129)	(0.0159)	(0.0288)	(0.0301)	(0.0081)	(0.0064)	(0.0092)	(0.0066)
	Diagnostics							
CIPS test	−4.1640***	−3.5140**	−3.2820*	−2.8880	−2.8870	−2.6110	−3.2570*	−3.2820*
△CIPS	−7.3720***	−6.3360***	−7.2030***	−6.6910***	−6.5790***	−5.6910***	−6.9410***	−5.5120***
CD tests	41.9030***	52.4970***	49.7440***	53.6010***	19.1330***	28.5200***	12.8180***	15.5700***
	Panel B: 1970–1989							
trade	−0.0003	−0.0094	−0.1286***	−0.0658	0.1451***	0.1271***	−0.1097**	−0.0232
	(0.0472)	(0.0462)	(0.0496)	(0.4479)	(0.0400)	(0.0330)	(0.0526)	(0.0415)
fopen	0.0849***	0.0870***	0.1265**	0.1034*	−0.0654***	−0.0306***	0.0487*	0.0285
	(0.0313)	(0.0319)	(0.0606)	(0.0586)	(0.0241)	(0.0099)	(0.0256)	(0.0205)
	Diagnostics							
CIPS test	−2.2260	−1.8510	−2.2400	−1.8080	−2.3020	−2.1120	−2.7890	−2.3990
△CIPS	−4.8700***	−4.1980**	−4.2070**	−3.7400**	−4.1890**	−3.7600**	−5.2080***	−3.5020**
CD tests	30.0900***	26.6130***	33.6460***	37.1020***	12.0930***	25.8120***	16.2940***	15.820***
	Panel C: 1970–2007							
trade	−0.0399	−0.0436	0.1619**	0.2432***	−0.0577	−0.1403**	0.0211	0.0564***
	(0.0509)	(0.0485)	(0.0769)	(0.0631)	(0.0391)	(0.0633)	(0.0328)	(0.0136)
fopen	−0.1453***	−0.1117*	0.1295***	0.1360***	−0.0203	0.0072	−0.0332*	−0.0137*
	(0.0547)	(0.0581)	(0.0346)	(0.0295)	(0.0138)	(0.0172)	(0.0197)	(0.0078)
	Diagnostics							
CIPS test	−3.0100*	−2.7260	−3.2260*	−2.9370	−2.7940	−1.9080	−2.6210	−1.8280
△CIPS	−4.9760***	−4.2520**	−5.0500***	−5.0100***	−4.3850***	−3.5620**	−4.0280**	−3.6500**
CD tests	13.2640***	25.6080***	45.9610***	48.1170***	18.2220***	20.2420***	11.8810***	18.8650***

Note: The values in parentheses are the standard errors based on a Newey–West type variance estimator.
***,**,* Significant at the 1%, 5%, and 10% levels, respectively.