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Discussion paper

GLOBALIZATION AND THE OUTPUT-INFLATION TRADEOFF: NEW TIME SERIES EVIDENCE

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Globalization and the Output-inflation Tradeoff: New Time Series Evidence*

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Abstract

Recent cross-country studies on the globalization and output-inflation tradeoff correlation find openness has no significant effect on OECD countries. Those studies assume parameter constancy across countries. In this paper, we argue that this assumption does not hold for major industrialized countries. Using individual time series analysis, we find the effect of openness on the output-inflation trade off differ in sign and size across countries. In contrast to previous cross-country studies, we find globalization has significantly changed some major industrialized countries' output inflation tradeoff. This has important implications for future theoretical and empirical research.

JEL classification: E31 E58 F10 F30 F41

Keywords: openness, output-inflation tradeoff, Phillips curve, time inconsistency theory

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1. Introduction

Recently there's a debate among leading economists on whether globalization has reduced long run inflation rate (see among others Rogoff, 2003; Ball, 2006). One key issue involved in this debate is whether globalization has changed the output-inflation tradeoff (or the slope of the Phillips curve) in OECD countries. If openness has indeed significantly affected the output-inflation tradeoff in OECD countries, the time inconsistency models predict that there should be a significant negative correlation between openness and the long run inflation rate. According to the most recent empirical studies, openness has no significant effect on the output-inflation tradeoff in OECD countries. This implies that the time inconsistency models on the globalization-inflation correlation do not apply in the OECD countries. However, we should not make such a conclusion too fast. In this paper, we show that the dominating empirical methodology in the studies on openness and the output-inflation tradeoff is problematic. Using an alternative empirical methodology we find that openness has significantly changed the output-inflation tradeoff in at least some major industrialized countries, but the sizes and directions of the effects differ across countries. This has important implications on future theoretical and empirical research on this topic.

The paper proceeds as follows: Section 2 briefly discusses the theoretical background. Section 3 introduces the empirical methodology. Section 4-5 present and discuss the empirical results. Section 6 concludes.

2. Globalization, the output-inflation tradeoff and long run inflation rate: theoretical background

In his seminal paper, Romer (1993) argued that domestic output expansion depreciates the real exchange rate in an open economy. The depreciation of real exchange rate makes output expansion more inflationary in an open economy than in a closed economy. In other words, the slope of the Phillips curve is steeper in an open economy than in a closed economy. This steepening effect is stronger when the degree of openness is higher. In this case, expanding domestic output level by an inflationary policy is less attractive for the central bank when the economy is more open. According to the time inconsistency theory, this leads to a lower long run inflation rate in a more open economy.

One weakness in Romer (1993)'s argument is that many small open economies can not affect its terms of trade. Lane (1997) provided another theoretical channel through which openness can affect the slope of the Phillips curve. He argued that for a small open economy, the traded sector faces perfect competition in the world market while non-traded sector faces imperfect competition and nominal rigidity. Imperfect competition in the non-traded sector means production in this sector is inefficiently low. Hence it's welfare-improving for the central banks to use inflationary policy to expand output in the non-traded sector. When the country becomes more open the relative size of non-traded sector declines. This means that inflationary policy has weaker effect on domestic output level and consumption welfare in a more open economy. Hence the central bank has smaller inflation bias in a more open economy. This leads to a lower long run inflation rate in a more open economy.

In contrast to Romer (1993) and Lane (1997), Razin and Loungani (2007) [1]argued that globalization weakens the link between domestic consumption and domestic production. Therefore, the Phillips curve is flatter in a more open economy. Because the central bank maximizes consumption welfare globalization also lowers its inflation bias and leads to a lower long run inflation rate.

A popular approach to test the globalization-Phillips curve correlation predicted in those models is cross-country regression (Temple, 2002; Daniels *et al*, 2005; Daniels and VanHouse, 2008; Badinger, 2009). And the dominating empirical result from those studies is that openness had no significant effect on the slope of the Phillips curve in the OECD. The only exception is the study by Daniels *et al* (2005). They found openness significantly flattened the Phillips curve in the OECD. However, Daniels and VanHouse (2008) found that the estimated effect turned insignificant once exchange rate passthrough was controlled for. The zero effect found by those studies leads the authors to conclude that the time inconsistency theory is not a satisfactory explanation for the openness-inflation correlation in the OECD (e.g. Temple, 2002; Badinger, 2009). However, in this paper, we argue that the zero effect found by those studies is a result of inappropriate empirical methodology. One should not draw the conclusion too fast.

The underlying key assumption of the cross-country regression is that the parameter of interest is constant across countries. This assumption is rather restrictive, because theoretically the sign of the effect of openness on the output-inflation tradeoff can be ambiguous. Using a new Keynesian model, Gali and Monacelli (2005) showed that the Phillips curve in an open economy is isomorphic to its closed economy

[1] Daniels and VanHouse (2006) also established a model to argue that openness makes the output-inflation tradeoff larger. However, their model is in fact about the tradeoff between output and the price level.

version if the preference is of the Dixit-Stiglitz type and exchange rate passthrough is complete. Specifically, domestic price inflation is described by the following equation:

$$\pi_t^d = \beta E_t \left\{ \pi_{t+1}^d \right\} + \lambda \widehat{rmc}_t \quad (1)$$

where $\lambda = \lambda(\beta, \theta)$, θ is the measure of price stickiness, rmc_t is the real marginal cost[2]. When exchange rate passthrough is complete, equation (1) is equivalent to:

$$\pi_t^d = \beta E_t \left\{ \pi_{t+1}^d \right\} + \delta_2 \widehat{y_t} \quad (2)$$

where $\widehat{y_t}$ is domestic output gap, $\delta_2 = \lambda \kappa_2$, $\kappa_2 = \kappa_2(\alpha, \Gamma_2)$, α is the degree of openness, Γ_2 is a vector of other structural parameters in the model[3].

From our definition of δ_2, we can write it as a function of the model's parameters. More specifically, $\delta_2 = \delta_2(\beta, \theta, \alpha, \Gamma_2)$. Obviously, the degree of openness α has an effect on the output-inflation tradeoff δ_2. From Gali and Monacelli (2005), it's easy to see that κ_2 is a quadratic equation of α, so only for certain values of parameters in Γ_2 $\frac{\partial \delta_2}{\partial \alpha}$ can have an unambiguous sign.

The point that the sign of $\frac{\partial \delta_2}{\partial \alpha}$ is ambiguous can be further strengthened by introducing strategic interactions between firms into the model. Sbordone (2008) showed that this can be done by substituting the assumption of Dixit-Stiglitz preferences by Kimball preferences. More specifically, Sbordone (2008) showed that with Kimball preference the open economy Phillips curve is of the following form:

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[2] Throughout, we will use $\hat{u}_t$ to denote the deviation of the variable u_t from its steady state level.

[3] Monacelli (2005) showed that when exchange rate passthrough is not complete, $\pi_t^d = \beta E_t \left\{ \pi_{t+1}^d \right\} + \delta_2 \widehat{y_t} + \delta_3 s_t$, where s_t is the supply shock. Since the supply shock term does not matter for our discussion on $\frac{\partial \delta_2}{\partial \alpha}$, we omit it in this section for simplicity.

$$\pi_t^d = \beta E_t\left\{\pi_{t+1}^d\right\} + \lambda_\alpha \widehat{rmc}_t \quad (3)$$

where $\lambda_\alpha = \lambda * \Omega(\alpha)$ and the sign of $\frac{\partial \Omega}{\partial \alpha}$ is ambiguous, depending on a number of structural parameters. Actually, Ω reflects the strategic interaction between firms, so we can expect that anything that affects the firms' competition environment is able to affect Ω.

Another source of ambiguity comes from the fact that openness has an ambiguous effect on price rigidity. Rogoff (2003) argued that globalization-induced competition could increase the price flexibility. That means $\frac{\partial \theta}{\partial \alpha} < 0$. However, as argued by Sbordone (2008), if globalization reduces trend inflation[4] firms will have less incentive to adjust their prices, which means price stickiness can rise with the degree of openness. Therefore, the sign of $\frac{\partial \theta}{\partial \alpha}$ is ambiguous. Since the output-inflation tradeoff δ_2 is a function of θ, the ambiguous sign of $\frac{\partial \theta}{\partial \alpha}$ can lead to an ambiguous sign of $\frac{\partial \delta_2}{\partial \alpha}$.

3. Our alternative empirical methodology

3.1. The benchmark empirical models

When the sign of $\frac{\partial \delta_2}{\partial \alpha}$ differs across countries, cross country regression imposes a false restriction on the model's parameter. This will cause serious estimation bias. To avoid this problem, we suggest using individual time series models to test whether globalization has affected OECD countries' output-inflation tradeoff[5]. More specifically, we take the following two specifications as our benchmark models for the test.

[4] Romer (1993), Lane (1997), Daniels and VanHouse (2006) provided theoretical arguments for this possibility.
[5] IMF (2006) used heterogeneous panel regression to perform the test. Although it allows for parameter heterogeneity in the regression, the main focus of the study is the cross-country mean effect. When the parameter of interest is not uniform across countries it's more interesting to know the country-specific effects from a policy perspective. Therefore we focus on individual time series studies of representative OECD countries rather than a panel study of the cross country mean effect.

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \varepsilon_{1t} \quad (4)$$

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha_{t-1} \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varepsilon_{2t} \quad (5)$$

where π_t^c is CPI inflation rate, $\overline{\pi_t^c}$ is its steady state level and $\widehat{\pi_t^c} = \pi_t^c - \overline{\pi_t^c}$. We use CPI inflation rather than domestic price inflation as the dependent variable because CPI inflation is more closely related to domestic welfare and central banks' target (Monacelli, 2005). Equation (4) is consistent with the backward-looking model of Ihrig *et al* (2007) while Equation (5) is consistent with the specification of Borio and Filardo (2007). Lee and Nelson (2007) showed that (5) is equivalent to the standard forward-looking new Keynesian Phillips curve (NKPC) but has the advantage to avoid parameter sensitivity to expectation horizon in the standard NKPC. Here we choose not to estimate a hybrid model with both forward-looking and backward-looking terms for two reasons: first, most empirical studies on the Phillips curve find one of those two terms dominates; second, previous studies reveal that when the true model is not a hybrid model, including both terms in the estimation will seriously bias the estimates of the model (Rudd and Whelan, 2005; Borio and Filardo, 2007). By using lagged output gaps on the right hand size, the Borio and Filardo (2007) specification has the additional advantage to reduce potential simultaneity bias caused by reversed causality from inflation to the output gap.

An important difference between our specifications here and the original specifications of Ihrig *et al* (2007) and Borio and Filardo (2007) is that we control for the effect of trend inflation on the slope of the Phillips curve. This is a necessary feature for constructing a valid test for the time inconsistency models on globalization and inflation. The time inconsistency models predict that openness can affect trend inflation rate by changing the output-inflation tradeoff. However, the causality can go the other way around. The state-dependent pricing models (Ball *et al*, 1988; Bakhshi *et al*, 2007) predict that trend inflation rate affects the output-inflation tradeoff. For this reason, if openness can affect trend inflation through other channels than affecting the output-inflation tradeoff[6], a significant effect of openness on the output-inflation tradeoff can be taken as evidence for the causality chain "openness→trend inflation→output inflation tradeoff" rather than

[6] One such other channel is suggested by Temple (2002). Temple (2002) argued that openness can affect the trend inflation through its effect on the exchange rate variability.

“openness→output inflation tradeoff→trend inflation”. Therefore, in order to lend support to the time inconsistency models one has to prove that openness can affect the output-inflation tradeoff besides its direct effect on trend inflation rate. Potential reversed causality from trend inflation to the output-inflation tradeoff also suggests a potential simultaneity bias in the OLS estimation. To avoid such a problem, we lag openness and trend inflation on the right hand side by one period in our specifications.

3.2. Potential omitted variable bias and the time-varying geographic instrument

Lagging variables on the right hand side can help avoid potential endogeneity bias caused by reversed causality, but it will not work when the endogeneity bias is caused by omitted variables. Our theoretical discussion in section 2 suggests that anything that can affect the industry structure and pricing behavior can affect the slope of the Phillips curve. When such factors are correlated to openness but omitted from our specifications, the OLS estimation will give biased results. Badinger (2009) suggested constructing an instrument variable for openness by geographic gravity models. In a cross-county setting, there’s danger that such an instrument is not exogenous. It is well known in the trade literature that the gravity models also have strong explanatory power for domestic trade pattern. That means geography can have direct effects on a country’s domestic industry structure and pricing behaviors beside its effect on the country’s international trade. In this case, the geographic instrument is not exogenous. However, if we can construct a time-varying instrument from the geographic variables the correlation between the instrument and geography in the error term will disappear, because there is almost no variation in geographic determinants of the output-inflation tradeoff within a country. In order to construct such a time-varying instrument, we estimate the following gravity model for each year:

$$\begin{aligned}&\log(Trade_{ijt} / GDP_{ijt}) = \delta_{0t} + \delta_{1t}\log(dist_{ij}) + \delta_{2t}\log(area_i * area_j)\\ &+\delta_{3t}Comlang_{ij} + \delta_{4t}Comborder_{ij} + \delta_{5t}Landlocked_{ij} + \omega_t\end{aligned} \quad (6)$$

where $dist_{ij}$ is the bilateral distance between two trade partners i and j, area is land area, Comlang is a dummy for common language, Comborder is a dummy for common land border, Landlocked is the number of landlocked countries in the two trade partners, ω_t is the error term. The instrument for openness is constructed as the sum of exponents of the fitted values across trading partners for each country. Because all the independent variables in the gravity model above are time-invariant, the time variation of the instrument comes from the changes in the parameters. Since the gravity model is estimated with parameter constancy assumption across trade partners, changes in the parameters are by construction global changes[7]. Therefore our instrument is exogenous if those global changes are exogenous. In section 5, we will formally control for the potential endogeneity of the instrument.

3.3. Measures of openness and data

Although our theoretical discussion in section 2 focused on trade openness, there are theories and evidence that financial openness can also affect a country's output-inflation tradeoff[8]. To test whether globalization in the broader sense has an effect on national output-inflation tradeoff, we use both trade and financial openness measures in our estimation[9]. Following Badinger (2009) we directly use the real trade openness measure from the Penn World Table 6.2. Financial openness is measured as total foreign assets and liabilities divided by GDP. We calculate it on the basis of the dataset of Lane and Milesi-Feretti (2006). Since the dataset of Lane and Milesi-Feretti (2006) only covers 1970-2000, the sample period for regressions with financial openness also covers only 1970-2000. Consumer Price inflation rates and real GDP (constant US dollar) data (1961-2007) are from the World Development Indicator database. We proxy the steady state inflation rates by the H-P filtered inflation rates (lambda=100). Bilateral trade and geographic data used to construct the instrument are from the dataset of Glick and Rose (2002) which covers 217 countries from 1948 to 1997. For this reason, sample end for the IV regressions is 1997 rather than 2000. The output gaps are estimated by the H-P filter with lambda=100.

4. Benchmark model results

4.1. The backward-looking model results

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[7] For example, Feyrer (2009) explained the time-varying property of the coefficients of bilateral distance in the gravity model by global changes in relative importance of air transportation with respect to sea transportation.

[8] Loungani *et al* (2001); Razin and Yuen (2002); Razin and Loungani (2007); Badinger (2009)

[9] Badinger (2009) suggested including both measures of openness in the estimation equation. However, this will cause serious collinearity problem. Badinger solved the problem by restricting the coefficient of financial openness to be equal to the coefficient of trade openness in his cross-country study. This restriction does not hold in our time series setting. Hence we use trade and financial openness as alternative measures of globalization in our tests rather than put them together in the test equations.

Table 1 presents the OLS estimation results of the benchmark backward-looking models. When the estimated coefficient of the interaction term between trend inflation and the output gap is not significant, we drop it from the specification and re-estimate the model. In this case, the results in Table 1 are those from the re-estimated models. The first observation from Table 1 is that the signs of both trade and financial openness differ across countries. Even when the signs are the same, the sizes of the effects are different. While globalization has no significant effect in several OECD countries, its effects are significant in some major industrialized countries (Trade openness is significant in France and Italy; Financial openness is significant in Italy, Netherlands and Switzerland). The obvious differences in signs and sizes of the effects across countries imply that the parameter constancy assumption of the cross country regression studies is unreliable.

Controlling omitted variable bias by instrument variable regressions (see results in Table 2) further strengthens the point that globalization matters in at least some major industrialized countries. Particularly, the Hausman tests reject the consistency of OLS estimator for trade openness in Canada and Switzerland, and eliminating potential omitted variable bias by IV regressions turns the coefficients of trade openness in those two countries from insignificant to significant The Hausman tests reject the consistency of OLS estimator for financial openness in more countries (Australia, Canada, France, and Sweden) and the coefficients of financial openness turn significantly negative in Australia and Canada when IV regressions are applied.

Combining the results in Table 1 and Table 2, the general finding of our benchmark backward-looking model is that trade openness significantly steepens the Phillips curve in France and Switzerland while it significantly flattens the Phillips curve in Canada and Italy; financial openness significantly steepens the Phillips curve in Italy while significantly flattens the Phillips curve in Australia, Canada, Netherlands and Switzerland. This sharply contrasts the cross-country regression result that openness has no effect on the Phillips curve in the OECD.

4.2. The forward-looking model results

Table 3 and 4 summarize the OLS and IV regression results from our benchmark forward-looking models. Similar to the backward-looking model results, the estimated coefficients of trade and financial openness differ in both sign and size. This again questions the reliability of cross-country regression results. The OLS regressions find a significant flattening effect of trade openness in Italy and UK. IV regression further finds a significant flattening effect of trade openness in Australia and this result is favored by the Hausman test.

As for financial openness, it has a significant steepening effect in Australia and a significant flattening effect in the UK according to the OLS regression results. Although the Hausman tests favors IV regressions for Canada and Sweden, the estimated coefficients of financial openness remain insignificant when estimated by IV regressions.

5. Robustness and the global inflation augmented models

The identifying assumption of our benchmark IV estimation is that global changes causing parameter variation in the gravity models are exogenous. In other words, those global changes should not be correlated with the model error terms. However, there is no guarantee that this is true. Ciccarelli and Mojon (2009) found that there's a global common factor in OECD countries' national inflation dynamics and they called this common factor "global inflation". If the global changes in the gravity model parameters are correlated to global inflation the benchmark models in section 4 will give us biased results. To check the robustness of our benchmark model results, we augment those benchmark models by the a proxy for "global inflation". More specifically, we estimate the following models:

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \phi_5 \pi_t^{c*} + \varepsilon_{3t} \quad (7)$$

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha_{t-1} \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varphi_4 \widehat{\pi_t^{c*}} + \varepsilon_{4t} \quad (8)$$

where the "global inflation" π_t^{c*} is defined as cross-country average of the domestic inflation rate of 22 OECD countries in Ciccarelli and Mojon (2009)s' sample. Ciccarelli and Mojon (2009) showed that this proxy is almost identical to the "global inflation" measures constructed by static and dynamic factor models.

5.1. The global inflation augmented backward-looking models

Before proceeds to the estimation results, further discussion on the stationarity of inflation rate in the backward-looking model is necessary. Estimated coefficients of lagged inflation in the benchmark backward-looking models are very close to 1 in most countries. That means national inflation rates may be non-stationary. Fanelli (2008) argued that although in theory inflation rate faced by the representative agents is stationary, aggregation of the data may still lead the aggregate inflation rate to be non-stationary. Table A1 reports the ADF and Phillips-Perron unit root test results of the national inflation rates in our sample. The tests fail to reject the unit root hypothesis at the 10% level for all the countries except Switzerland.

Culver and Papell (1997), Basher and Westerlund (2008) argued that the inability of individual unit root tests to reject the null hypothesis is due to the lack of power in finite sample. They proposed to use panel unit root tests to increase the power of the tests. They found inflation rate stationary by their panel unit root tests. However, the tests they applied rely on some restrictive assumptions. Particularly most of those tests assume no dynamic interdependencies and requires a large cross-section. Palm *et al* (2008) proposed a cross-sectional dependence robust block bootstrap panel unit root test (henceforth the RBB test) which is robust to very general error structures including the case with dynamic interdependencies. Their test is valid for finite N, which is also desirable for our purpose. When the cross-section is large and the null hypothesis is rejected, we only know that inflation rate is stationary in at least some countries, which is not very informative. As argued by Culver and Papell (1997), a rejection of the unit root null hypothesis is more in favor of the stationarity assumption of individual countries' inflation rate when the cross-section is smaller. Table A2 summarizes our panel unit root test results. Results in the upper panel of Table A2 are based on the panel unit root test of Pesaran (2007), which is also applied in Basher and Westerlund (2008). Consistent with the finding of Basher and Westerlund (2008), the Pesaran (2007) test rejects the unit root hypothesis. However, the more general RBB panel unit root test failed to reject the unit root hypothesis.

When inflation rate is an I(1) variable, our benchmark backward-looking model is not balanced since all the variables on the right hand side are I(0) variables. To balance the model we need some I(1) variables on the right hand side as additional independent variables. Unit root test results of the global inflation rate reported in Table A1 suggest that it is an I(1) variable. Rearranging equation (7) we can get the following equation:

$$\Delta\pi_t^c = \phi_0 + \vartheta(\pi_{t-1}^c + \phi_6\pi_{t-1}^{c\,*}) + \phi_2\widehat{y_t} + \phi_3\alpha_{t-1}\widehat{y_t} + \phi_4\overline{\pi_{t-1}^c}\widehat{y_t} + \phi_5\Delta\pi_t^{c*} + \varepsilon_{3t} \quad (9)$$

where $\vartheta = \phi_1 - 1, \phi_6 = \dfrac{\phi_5}{\phi_1 - 1}$. Hence if domestic and global inflation are cointegrated and we take $\mu' Y_{t-1} = \pi_{t-1}^c + \phi_5\pi_{t-1}^{c\,*}$ as a variable, all the variables in equation (9) are stationary and we can apply the standard inference to it.

We proceed in two steps. First we test for cointegration between domestic and global inflation based on a bivariate VAR of $Y_t = (\pi_t, \pi_t^*)'$. More specifically we estimate the cointegrating relation from the following error correction model with full information maximum likelihood method:

$$\Delta Y_t = \sigma\mu' Y_{t-1} + \sum_{i=1}^{p} \Delta Y_{t-i} + \xi_t$$

where ξ_t is the error term and μ is the cointegrating vector. Here we don't include an intercept term in the VAR since we expect no deterministic trend in both domestic and global inflation rate. The lag order p is selected according to the Schwartz information criterion. The test results are summarized in Table A3. Since we only find evidence for cointegration in 5 sample countries, we present the results for those five countries only.

Our second step is to substitute the estimated cointegrating relation (henceforth ECM) for $\mu' Y_{t-1}$ and estimate equation (9) with usual least square estimators. The results are summarized in Table 5 and Table 6. An important empirical issue is the exogeneity of global inflation in the model. Dees *et al* (2007) suggested that we can test the exogeneity of global inflation formally by testing the significance of the ECM term in the equation for global inflation in the bivariate VAR of $Y_t = (\pi_t, \pi_t^*)'$. We present the t test statistics in Table A3 and the results reveal that global inflation is exogenous for all the five countries.

The signs and sizes of the effects of trade and financial openness on the Phillips curve are still different across countries in the sample, which contrasts the parameter constancy assumption of the cross-country studies. The evidence in support of a significant effect of globalization is much weaker than that from the benchmark backward-looking model. However, the conclusion of no effect in major industrialized countries from the cross-country studies does not apply to the United States. Both OLS and IV regressions find a significant steeping effect of financial openness on the Phillips curve in the US. Although the OLS estimate of trade openness coefficient is not significant in the US, the IV estimate is negatively significant and favored by the Hausman test.

5.2. The global inflation augmented forward-looking models

Since stationarity is not a concern for the inflation gap we directly apply usual OLS and IV regression methods to the global inflation augmented forward-looking models. Table 7 and 8 present the results. Consistent with our benchmark models and the global inflation augmented backward-looking models, we find the signs and sizes of the estimated coefficients of trade and financial openness differ across countries. Compared to the results from the backward-looking models, the results from the forward-looking models are more supportive for the hypothesis that globalization has significantly changed some major industrialized countries' output-inflation tradeoff. The OLS regression results suggest that trade openness significantly flattens the Phillips curve in Canada while its effects are not significant in all other sample countries. The OLS results also suggest that financial openness significantly steepens the Phillips curve in Australia, France and US while it significantly flattens the Phillips curve in the Netherlands. The IV regression results are generally not favored by the Hausman tests except the case for trade openness in the Netherlands. However, the qualitative results from the IV estimation are not very different from the OLS results. Trade openness is still positive but insignificant in the Netherlands.

6. Conclusion

Recent cross-country regression studies find openness has no effect on industrialized countries' output-inflation tradeoff. The underlying assumption of those studies is parameter constancy across countries. In this paper we argue that the validity of this assumption is not guaranteed from a theoretical perspective. Our individual time series analysis reveals that the signs and sizes of the effects of openness on the slope of the Phillips curve differ across countries. This questions the reliability of the empirical results from the cross-country regressions. Our results suggest that trade and financial globalization have at least affected some major industrialized countries' output-inflation tradeoff, so globalization matters. However, these results are not sufficient to support the current time inconsistency models of globalization-inflation correlation. All the current theoretical models assume that the direction of openness' effect on the slope of the Phillips curve is one way and they all predict that the one-way effect on the Phillips curve finally reduces long run inflation rate. Since our results suggest that the effect of openness on the slope of the Phillips curve may differ in signs across countries, we have to observe that openness reduces long run inflation rate in some countries while increases long run inflation rate in other countries to reconcile the theory with the evidence. Since most previous studies on the openness-inflation correlation assume parameter constancy across countries, we cannot draw the conclusion based on those studies. Moreover, as suggested by Temple (2002) and Rogoff (2003), globalization may have affected long run inflation rate through other channels besides the Phillips curve channel. This further complicates the identification problem. We leave this to future research.

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Table 1 OLS estimation results of the benchmark backward-looking model

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \varepsilon_{1t}$$

	Measure of α : trade openness			Measure of α : financial openness		
	ϕ_3	ϕ_4	$\overline{R^2}$	ϕ_3	ϕ_4	$\overline{R^2}$
Australia	-11.8	-	0.76	-2.63	-	0.72
	(9.92)			(2.13)		
Canada	-2.35	-	0.81	-0.66	-	0.77
	(1.58)			(0.49)		
France	6.27***	0.32***	0.90	0.42	0.27***	0.92
	(2.01)	(0.07)		(0.29)	(0.07)	
Italy	-5.04*	0.09**	0.87	1.56*	0.25***	0.85
	(2.74)	(0.05)		(0.87)	(0.08)	
Netherlands	-2.08	-	0.67	-0.51**	-	0.80
	(2.08)			(0.22)		
Sweden	-8.23	-	0.62	-0.30	-	0.63
	(6.30)			(0.58)		
Switzerland	-2.18	-	0.74	-0.41***	-	0.79
	(2.54)			(0.14)		
UK	-1.35	-	0.65	-0.13	-	0.61
	(2.35)			(0.17)		
US	-0.28	0.02***	0.78	0.04	0.03***	0.74
	(0.28)	(0.01)		(0.05)	(0.01)	

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively.

Table 2 IV estimation results of the benchmark backward-looking model

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \varepsilon_{1t}$$

	Measure of α : trade openness				Measure of α : financial openness			
	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman
Australia	-25.2**	-	0.77	0.38	-9.03*	-	0.71	3.04***
	(12.7)				(4.74)			
Canada	-8.87**	-	0.79	2.15**	-5.50**	-	0.70	2.51**
	(3.59)				(2.40)			
France	4.14	0.29**	0.89	0.33	2.83	0.59	0.87	-2.46**
	(6.63)	(0.13)			(2.63)	(0.39)		
Italy	-6.76*	0.11**	0.86	-0.02	4.44	0.32	0.85	0.37
	(3.46)	(0.05)			(9.10)	(0.29)		
Netherlands	4.13	-	0.67	-1.44	0.65	-	0.78	-1.46
	(3.89)				(1.27)			
Sweden	-28.2	-1.25*	0.92	0.76	0.13	-	0.57	2.10**
	(24.9)	(0.69)			(2.54)			
Switzerland	11.9**	1.41**	0.73	-1.71*	-0.27	-	0.78	-0.44
	(6.00)	(0.59)			(0.32)			
UK	-3.76	-	0.63	0.78	-0.11	-	0.58	0.01
	(6.49)				(0.39)			
US	-1.50***	0.013**	0.78	1.10	-0.27***	-	0.73	0.96
	(0.53)	(0.005)			(0.06)			

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively. Hausman test statistics reported in the table are t values calculated following the two-step procedure of Davidson and Mackinnon (1989) and we base the calculation on Newey-West HAC standard errors.

Table 3 OLS estimation results of the benchmark forward-looking model

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varepsilon_{2t}$$

	Measure of α : trade openness			Measure of α : financial openness		
	φ_2	φ_3	$\overline{R^2}$	φ_2	φ_3	$\overline{R^2}$
Australia	3.28	0.62***	0.18	9.85*	1.90***	0.23
	(12.3)	(0.22)		(5.30)	(0.57)	
Canada	1.22	0.28**	0.36	-0.16	0.21**	0.36
	(2.18)	(0.11)		(0.89)	(0.10)	
France	1.18	0.21***	0.43	0.35	0.25***	0.51
	(1.78)	(0.04)		(0.49)	(0.07)	
Italy	-6.62**	-	0.17	1.54	0.24**	0.21
	(2.66)			(1.98)	(0.09)	
Netherlands	0.49	-	0.01	-0.33	-	0.09
	(2.61)			(0.34)		
Sweden	-10.2	-	0.22	-0.27	-	0.16
	(6.44)			(0.57)		
Switzerland	-2.63	-	0.64	-0.24	-	0.63
	(1.62)			(0.19)		
UK	-6.59***	-	0.31	-0.39***	-	0.31
	(1.64)			(0.11)		
US	0.10	0.004***	0.53	0.04	0.04**	0.53
	(0.54)	(0.001)		(0.10)	(0.02)	

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively.

Table 4 IV estimation results of the benchmark forward-looking model

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varepsilon_{2t}$$

	Measure of α : trade openness				Measure of α : financial openness			
	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman
Australia	-18.3*	0.53***	0.19	2.17**	-7.40**	-	0.18	1.50
	(9.81)	(0.16)			(2.71)			
Canada	-7.88**	-	0.27	1.37	7.91	0.90*	-0.32	-2.13**
	(3.62)				(6.47)	(0.46)		
France	-7.02	-	0.28	0.45	-3.07**	-	-0.34	0.91
	(5.78)							
Italy	-10.2**	-	0.16	1.02	-3.30	-	0.15	1.04
	(5.11)				(2.08)			
Netherlands	4.20	-	0.01	-1.56	0.63	-	0.07	-0.69
	(4.08)				(1.49)			
Sweden	-26.5**	-	0.17	1.61	-1.69	-	0.14	-2.16**
	(11.3)				(1.46)			
Switzerland	-1.37	-	0.64	-0.43	-0.17	-	0.64	0.02
	(2.63)				(0.22)			
UK	-8.51**	-	0.31	0.66	-0.49**	-	0.31	0.60
	(3.47)				(0.19)			
US	-2.34**	-	0.45	0.65	-0.39***	-	0.53	1.24
	(1.26)				(0.12)			

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively. Hausman test statistics reported in the table are t values calculated following the two-step procedure of Davidson and Mackinnon (1989) and we base the calculation on Newey-West HAC standard errors.

Table 5 OLS estimation results of the global inflation augmented backward-looking model

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \phi_5 \pi_t^{c*} + \varepsilon_{3t}$$

	Measure of α : trade openness			Measure of α : financial openness		
	ϕ_3	ϕ_4	$\overline{R^2}$	ϕ_3	ϕ_4	$\overline{R^2}$
Australia	-5.62	-	0.54	-0.04	-	0.56
	(9.08)			(1.45)		
Canada	-1.34	-	0.58	-0.26	-	0.60
	(1.03)			(0.31)		
Sweden	1.74	-	0.51	-0.02	-	0.47
	(4.51)			(0.78)		
UK	0.47	-	0.57	0.05	-	0.56
	(3.92)			(0.27)		
US	-0.25	0.017*	0.75	0.10***	0.043***	0.82
	(0.44)	(0.009)		(0.03)	(0.008)	

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively.

Table 6 IV estimation results of the global inflation augmented backward-looking model

$$\pi_t^c = \phi_0 + \phi_1 \pi_{t-1}^c + \phi_2 \widehat{y_t} + \phi_3 \alpha_{t-1} \widehat{y_t} + \phi_4 \overline{\pi_{t-1}^c} \widehat{y_t} + \phi_5 \pi_t^{c*} + \varepsilon_{3t}$$

	Measure of α : trade openness				Measure of α : financial openness			
	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman
Australia	-19.6*	-	0.54	0.82	-1.29	-	0.54	0.22
	(10.7)				(3.15)			
Canada	-4.77*	-	0.57	1.04	-1.84	-	0.59	0.62
	(2.87)				(1.82)			
Sweden	19.3	-	0.53	-0.42	4.12	-	0.52	-0.51
	(19.8)				(3.72)			
UK	-5.12	-	0.53	2.52**	-0.22	-	0.52	1.70*
	(6.37)				(0.37)			
US	-2.54**	-	0.76	2.13**	-0.28***	-	0.80	1.52
	(0.59)				(0.07)			

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively. Hausman test statistics reported in the table are t values calculated following the two-step procedure of Davidson and Mackinnon (1989) and we base the calculation on Newey-West HAC standard errors.

Table 7 OLS estimation results of the global inflation augmented forward-looking model

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha_{t-1} \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varphi_4 \widehat{\pi_t^{c*}} + \varepsilon_{4t}$$

	Measure of α : trade openness			Measure of α : financial openness		
	φ_2	φ_3	$\overline{R^2}$	φ_2	φ_3	$\overline{R^2}$
Australia	0.41	0.45**	0.43	7.21*	1.43***	0.46
	(8.86)	(0.17)		(3.98)	(0.52)	
Canada	-2.86***	-	0.66	-0.35	0.13*	0.70
	(0.60)			(0.57)	(0.07)	
France	3.43	0.15***	0.67	0.59**	0.18***	0.81
	(2.14)	(0.04)		(0.27)	(0.04)	
Italy	-2.45	-	0.81	1.55	0.12*	0.84
	(2.05)			(0.94)	(0.06)	
Netherlands	0.58	-	0.42	-0.75***	-0.62***	0.58
	(1.52)			(0.12)	(0.22)	
Sweden	-6.03	-	0.42	-0.06	-	0.54
	(4.47)			(0.47)		
Switzerland	-1.46	-	0.69	-0.11	-	0.69
	(2.04)			(0.22)		
UK	-2.23	-	0.60	-0.08	-	0.58
	(4.50)			(0.34)		
US	-0.37	0.03*	0.71	0.09*	0.05***	0.80
	(0.50)	(0.01)		(0.05)	(0.01)	

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively.

Table 8 IV estimation results of the global inflation augmented forward-looking model

$$\widehat{\pi_t^c} = \varphi_0 + \varphi_1 \hat{y}_{t-1} + \varphi_2 \alpha_{t-1} \hat{y}_{t-1} + \varphi_3 \overline{\pi_{t-1}^c} \hat{y}_{t-1} + \varphi_4 \widehat{\pi_t^{c*}} + \varepsilon_{4t}$$

	Measure of α : trade openness				Measure of α : financial openness			
	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman	ϕ_3	ϕ_4	$\overline{R^2}$	Hausman
Australia	-8.20	-	0.40	-1.21	-5.17***	-	0.43	0.23
	(9.00)				(1.66)			
Canada	-5.40***	-	0.63	0.43	2.97	0.41**	0.59	-1.42
	(2.12)				(2.69)	(0.18)		
France	-1.64	-	0.59	-0.49	-1.16	-	0.67	-0.30
	(5.69)				(1.51)			
Italy	-4.68*	-	0.82	0.07	-0.71	-	0.84	-1.54
	(2.45)				(0.98)			
Netherlands	3.80	-	0.42	-2.52**	1.65*	-	0.52	-1.19
	(2.43)				(0.94)			
Sweden	-9.51	-	-0.15	0.56	0.75	-	0.43	-0.61
	(12.8)				(1.92)			
Switzerland	0.20	-	0.69	-0.32	-0.16	-	0.69	0.33
	(2.61)				(0.26)			
UK	-6.93**	-	0.58	0.14	-0.37*	-	0.55	0.31
	(3.24)				(0.21)			
US	-3.05***	-	0.63	0.38	-0.14**	-	0.71	1.26
	(1.11)				(0.06)			

Note: Newey-West HAC standard errors in parentheses; ***, **, * denote significance at 1%, 5% and 10% level respectively. Hausman test statistics reported in the table are t values calculated following the two-step procedure of Davidson and Mackinnon (1989) and we base the calculation on Newey-West HAC standard errors.

Table A1 individual unit root tests for national and global inflation rates

ADF test: 1961-2000					
Country name	p	Test statistics	1% critical value	5% critical Value	10% critical value
Australia	0	-1.74	-3.61	-2.94	-2.61
Canada	0	-1.69	-3.61	-2.94	-2.61
Switzerland	1	-3.39**	-3.61	-2.94	-2.61
France	0	-1.22	-3.61	-2.94	-2.61
Italy	0	-1.56	-3.61	-2.94	-2.61
Netherlands	0	-1.95	-3.61	-2.94	-2.61
Sweden	0	-1.99	-3.61	-2.94	-2.61
UK	0	-2.18	-3.61	-2.94	-2.61
US	2	-1.83	-3.61	-2.94	-2.61
Global inflation	1	-1.85	-3.61	-2.94	-2.61

Phillips-Perron test: 1961-2000					
Country name	p	Test statistics	1% critical value	5% critical Value	10% critical value
Australia	0	-1.89	-3.61	-2.94	-2.61
Canada	0	-1.70	-3.61	-2.94	-2.61
Switzerland	0	-2.79*	-3.61	-2.94	-2.61
France	0	-1.22	-3.61	-2.94	-2.61
Italy	0	-1.74	-3.61	-2.94	-2.61
Netherlands	0	-1.95	-3.61	-2.94	-2.61
Sweden	0	-1.99	-3.61	-2.94	-2.61
UK	0	-2.24	-3.61	-2.94	-2.61
US	0	-1.98	-3.61	-2.94	-2.61
Global inflation	0	-1.54	-3.61	-2.94	-2.61

Notes: Calculation by Eviews 6.0.

ADF test regression: $\Delta\pi^c_t = a_0 + a_1\pi^c_{t-1} + \sum_{i=1}^{p} b_i \Delta\pi^c_{t-i} + e^1{}_t$, where p is the lag order selected by the Schwartz information criterion. Null hypothesis: $a_1 = 0$.

Phillips-Perron test regression: $\Delta\pi^c_t = c_0 + c_1\pi^c_{t-1} + e^2{}_t$. Null hypothesis: $c_1 = 0$. Adjusted t test statistics calculated with Newey-West bandwidth using Bartlett kernel.

***, **, * denote significance at 1%, 5% and 10% level respectively.

Table A2 Panel Unit root tests for national inflation rates

Pesaran (2007) truncated CIPS panel unit root test			Sample period: 1961-2000
Test statistics	1% critical value	5% critical Value	10% critical value
-3.46***	-2.55	-2.33	-2.21

RBB group mean panel unit root test			Sample period: 1961-2000
Test statistics	1% critical value	5% critical Value	10% critical value
-6.84	-8.71	-7.61	-6.92

Notes: Calculation by Stata programming.

Pesaran test regression: $\Delta\pi_{it}^{c} = \lambda_i + \rho_i \pi_{i,t-1}^{c} + \eta_i \overline{\pi_{t-1}^{cm}} + \sum_{j=0}^{p} \phi_{ij} \overline{\Delta\pi_{t-i}^{cm}} + \sum_{j=1}^{p} \varphi_{ij} \Delta\pi_{i,t-i}^{c} + e^{3}{}_{it}$, where $\overline{\pi_{t-1}^{cm}}$ is the cross-sectional mean of the inflation rate. Null hypothesis: $\rho_i = 0$ for all *i*. Alternative hypothesis: $\rho_i < 0$, *i*=1, 2, …, N_1, $\rho_i = 0$, *i*=N_1+1, N_1+2, …, N.

RBB test statistics calculated as the cross-sectional mean of T times the individual regression coefficients of the following equation: $\Delta\pi_{it}^{c} = d_i \pi_{i,t-1}^{c} + e^{4}{}_{it}$. The bootstrap critical values are obtained on the basis of 2000 simulations. The block length $b = 1.75 * T^{1/3}$.

***, **, * denote significance at 1%, 5% and 10% level respectively.

Table A3 Cointegration tests

Australia			
(1) Trace test			
Null hypothesis	Trace statistics	0.05 critical value	Probability
r=0	11.50*	12.32	0.069
r≤1	0.39	4.13	0.594
(2) Maximum Eigenvalue test			
Null hypothesis	Maximum Eigenvalue Statistics	0.05 critical value	Probability
r=0	11.10*	11.22	0.053
r≤1	0.39	4.13	0.594
Estimated cointegrating relation		$\hat{\beta}' Y_{t-1} = \pi_{t-1} - \underset{(0.08)}{1.00} \pi_{t-1}^{*}$	

t statistics of the error correction term in the equation of global inflation : -0.20

Canada			
(1) Trace test			
Null hypothesis	Trace statistics	0.05 critical value	Probability
r=0	14.97**	12.32	0.018
r≤1	0.37	4.13	0.606
(2) Maximum Eigenvalue test			
Null hypothesis	Maximum Eigenvalue Statistics	0.05 critical value	Probability
r=0	14.60**	11.22	0.012
r≤1	0.37	4.13	0.606
Estimated cointegrating relation		$\hat{\beta}' Y_{t-1} = \pi_{t-1} - \underset{(0.04)}{0.81} \pi_{t-1}^{*}$	

t statistics of the error correction term in the equation of global inflation : 0.60

UK
(1) Trace test

Null hypothesis	Trace statistics	0.05 critical value	Probability
r=0	14.04**	12.32	0.026
r≤1	0.38	4.13	0.600

(2) Maximum Eigenvalue test

Null hypothesis	Maximum Eigenvalue Statistics	0.05 critical value	Probability
r=0	13.66**	11.22	0.018
r≤1	0.38	4.13	0.600

Estimated cointegrating relation $\hat{\beta}' Y_{t-1} = \pi_{t-1} - \underset{(0.09)}{1.18}\pi_{t-1}^{*}$

t statistics of the error correction term in the equation of global inflation : -0.10

Sweden

(1) Trace test

Null hypothesis	Trace statistics	0.05 critical value	Probability
r=0	18.02***	12.32	0.005
r≤1	0.39	4.13	0.597

(2) Maximum Eigenvalue test

Null hypothesis	Maximum Eigenvalue Statistics	0.05 critical value	Probability
r=0	17.63***	11.22	0.003
r≤1	0.39	4.13	0.597

Estimated cointegrating relation $\hat{\beta}' Y_{t-1} = \pi_{t} - \underset{(0.06)}{0.95}\pi_{t-1}^{*}$

t statistics of the error correction term in the equation of global inflation : -0.79

US

(1) Trace test

Null hypothesis	Trace statistics	0.05 critical value	Probability
r=0	21.87***	12.32	0.001

r≤1	0.54	4.13	0.526

(2) Maximum Eigenvalue test

Null hypothesis	Maximum Eigenvalue Statistics	0.05 critical value	Probability
r=0	21.33***	11.22	0.001
r≤1	0.54	4.13	0.526

Estimated cointegrating relation $\hat{\beta}' Y_{t-1} = \pi_{t-1} - \underset{(0.04)}{0.75} \pi_{t-1}^*$

t statistics of the error correction term in the equation of global inflation : 0.46