Openness and the Output-Inflation Tradeoff*

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Abstract

Standard open economy models predict that openness to trade should exert a positive effect on the slope of the output-inflation tradeoff, or Phillips curve, but such a proposition finds very little support in the existing empirical literature. We propose a new test of this hypothesis based on new measures of the slope of the Phillips curve and more general cross-country regression models. The results provide strong empirical support for the standard theoretical prediction.

KEYWORDS: Openness, Inflation, Phillips curve.
JEL CLASSIFICATION: E31, E32, F41.

1 Introduction

This paper uses cross-sectional data on 20 countries to test the hypothesis that the slope of the short-run Phillips curve (drawn in output-inflation space) varies positively with openness to trade. We present a series of regressions in which the slope of the Phillips curve is the dependent variable and the regressor set comprises a number of controls suggested by both closed and open economy models. The principal finding is that a country's openness to trade exerts a positive and robust effect on the slope of its Phillips curve (or output-inflation tradeoff) provided that one controls for a country's exchange rate regime.

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This result contrasts with recent findings in Temple (2002), which indicate that openness exerts no systematic effect on the slope of the Phillips curve. We suggest two reasons for the differences between past results and our own. First, previous research has used a measure of the slope of the Phillips curve due to Ball, Mankiw and Romer (1988).[1] This index is derived from very parsimonious regression models which fail to control for a variety of long-run influences on output and inflation and therefore lead to biased measures of the slope of the Phillips curve, a point made in the commentary on BMR. To deal with this problem we replace the BMR index with an alternative measure of the slope of the Phillips curve based on the results in Bowdler (2002).

Second, previous studies test for a linear effect of openness on the terms of the output-inflation tradeoff, while economic theory predicts a relationship featuring openness plus its interaction with a country's contribution to world GDP, its chosen exchange rate regime and its historical inflation experience.[2] The second key innovation of this paper is to condition the analysis on those interaction terms. The interaction between openness and the exchange rate regime is the empirically relevant effect that has been overlooked by past research. Conditioning on that interaction term establishes a strong and stable relationship between openness and the slope of the Phillips curve.

The research described in this paper is interesting for a number of reasons. As we discuss below, a positive relationship between openness and the slope of the Phillips curve is a central prediction of a range of macroeconomic models, e.g. those due to Romer (1993) and Lane (1997). Temple (2002) notes that the lack of empirical support for such models that has been recorded up to now represents an important puzzle in open economy macroeconomics. A basic contribution of this paper is to reconcile theoretical and empirical research in this field. Further, the model that we estimate pins down the channels through which openness affects the slope of the Phillips curve, allowing us to discriminate between rival theories. We find that Lane's (1997) model featuring small economies and exchange rates determined as in the Mundell-Fleming model receives much stronger support than Romer's (1993) model of large economies, each capable of influencing the international price of goods.

The remainder of the paper expands on these points and has the following structure. Section 2 discusses the theoretical models. Section 3 reviews some empirical tests of those models, focusing on issues of measurement and model specification. Section 4 describes data collection and Section 5 reports new empirical results. Section 6 investigates the robustness of our core findings and Section 7 rounds off with a summary.

[1] Hereafter, BMR.

[2] The models underpinning these predictions are described in Section 2 of the paper.

2 Theoretical Perspectives

In this section we first examine closed economy models of the output-inflation tradeoff proposed during the 1970s and 1980s. Open economy models are then discussed, with particular attention paid to the conditions necessary for openness to affect the slope of the Phillips curve.

Closed Economy Models The modern literature on the output-inflation tradeoff begins with the work of Lucas (1972, 1973). According to this approach agents face a signal extraction problem following unanticipated shocks to the money supply, due to the fact that they cannot observe the current general level of prices. As is well known, this signal extraction problem is solved using information on the volatilities of general prices and relative prices, such that those magnitudes affect the slope of the short-run Phillips curve. Formally, the model yields the following hypothesis:

> Hypothesis One: If the ratio of volatility in the general price level to volatility in idiosyncratic prices is high then the Phillips curve will be steeply sloped in output-inflation space, while if the ratio is small then the Phillips curve will be shallow in output-inflation space.

An alternative theory of the output-inflation tradeoff is provided by BMR. According to their model firms plan to set prices infrequently due to the presence of 'menu costs'. Further, the planned duration of the period of price inertia will be negatively related to mean inflation, for high inflation erodes a firm's relative price and increases the incentive to pay the menu cost more often. As shorter periods of price inertia lead to aggregate demand feeding into higher prices and inflation more rapidly, the short-run Phillips curve will be steeper in output-inflation space the higher mean inflation. This reasoning is summarised in the following hypothesis:

> Hypothesis Two: A low mean inflation rate will lead to a shallow output-inflation tradeoff, whilst a high mean inflation rate will lead to a steep output-inflation tradeoff.

As hypotheses one and two make distinct predictions concerning the slope of the Phillips curve, they can be used to test between the Lucas and BMR models.[3]

[3]The BMR model also generates the result that the output-inflation tradeoff is more shallow the smaller the volatility of the unexpected component of the price level, see BMR (1988). However, it is the role of inflation that BMR emphasise in their paper.

Open Economy Models Open economy effects on the slope of the output-inflation tradeoff have been analysed by Romer (1993), who examined the case of an economy large enough to influence the international price of goods. In such an economy, prices for domestically produced output are controlled by two types of firm. The first type can adjust prices at any point in time, while the second can only adjust them infrequently.[4] In contrast, the price of imported goods can adjust freely at all times and is determined as the foreign price of goods multiplied by the nominal exchange rate.

An expansion of aggregate demand will raise import expenditures in such an economy, and will also lead to extra domestically produced output being supplied to the world market. Now, given that the country in question is large, the exchange rate will depreciate in order to clear the world market for tradables. This forces up import prices and thereby raises consumer prices. Additionally, if imports are used as inputs to domestic production then home producer prices will rise for cost-push reasons. Both effects will increase the amount of inflation associated with a given expansion of aggregate demand, i.e. the short-run Phillips curve will be steeper than in the closed economy case.[5] Further, the steepening of the Phillips curve will be more pronounced in relatively open economies because a country's openness determines the amount of inflation that it imports following a depreciation.

The Romer prediction only applies, however, to countries large enough to influence world prices. Small open economies typically represent a tiny fraction of world trade and therefore cannot induce systematic fluctuations in the real exchange rate through their impact on the international price of goods.[6] In such cases the openness of the economy holds no implications for the slope of the short-run Phillips curve. Hence, Romer's result is a conditional one, which we summarise in the following hypothesis:

> Hypothesis Three: The slope of the Phillips curve will respond positively to openness, but the effect will be weaker in relatively small economies.

Lane (1997) modifies Romer's model so that exchange rates are determined as in the Mundell-Fleming model. As is well known, this model generates the result that

[4]The reason for this inertia is not spelt out by Romer, but potential reasons are provided by the closed economy models described above.

[5]The output expansion will be the same in the closed and open economy cases, for although greater openness leads to greater 'demand leakages' via import spending, it also implies greater increases in exports following the depreciation. Essentially, demand innovations affect output in the tradable and non-tradable sectors symmetrically, while they impact prices asymmetrically (due to the role of the exchange rate in setting import prices), hence the Phillips curve is steeper in an open economy.

[6]This may not be true in the case of a small country that makes a relatively large contribution to world production of a highly specialised good. However, in practice most countries's trade is very diversified, so it is absolute size that determines their ability to influence world prices.

monetary policy expansions induce depreciation of the nominal exchange rate, while contractions induce appreciations. Of course, these are precisely the exchange rate dynamics which ensure that any increment to detrended GDP raises inflation in an open economy by more than it would in a closed economy. Further, a more open economy will face a steeper Phillips curve irrespective of whether or not it is large enough to influence the international price of goods. The Lane model does, however, predict an interaction between openness and a country's exchange rate regime. This prediction is summarised in the following hypothesis:

> Hypothesis Four: Greater openness will increase the slope of the Phillips curve, but the effect will grow weaker and weaker as the monetary policy authority increases its commitment to fixing the exchange rate.

Taylor (2000) suggests a further hypothesis pertaining to the relationship between openness and the slope of the output-inflation tradeoff. This model suggests that the amount of pass-through from exchange rate driven import price shocks to consumer prices will be smaller in economies that have experienced low inflation in the past.[7] In such countries firms reason that the exchange rate depreciations underpinning increases in consumer prices will quickly be reversed through a tightening of monetary policy. Consequently they will choose to limit the extent to which exchange rate induced import price shocks are passed through to consumer prices. Embedding this idea in the Lane model yields our final hypothesis:

> Hypothesis Five: Openness will not exert such a powerful effect on the slope of the output-inflation tradeoff in economies that have a history of low inflation.

3 Testing the Theoretical Models

A number of studies report empirical tests of the hypotheses described in Section 2. For example, BMR (1988) find that mean inflation helps to explain the slope of the Phillips curve while the volatility of nominal GDP growth (a measure of the severity of the Lucas signal extraction problem) does not. More recently, Temple (2002) extends the BMR analysis to examine the impact of openness. The results not only indicate that openness exerts an insignificant effect on the slope of the Phillips curve, but that the estimated relationship is incorrectly signed. The robustness of this finding is confirmed using a least trimmed squares estimator and when measuring the slope of

[7] This is the implication that Choudhri and Hakura (2001) draw from Taylor's model, and for which they find strong empirical support.

the Phillips curve using the sacrifice ratios calculated by Ball (1994) and the benefit ratios calculated by Jordan (1997).

The Temple finding represents something of a puzzle, for the models that predict a positive relationship between openness and the slope of the Phillips curve are built from quite plausible foundations. The key assumptions are that monetary policy affects output in both the tradable and non-tradable sectors of the economy, that monetary policy expansions depreciate the exchange rate and contractions appreciate the exchange rate, and that exchange rate driven fluctuations in import prices are passed through to consumer prices. Empirical evidence supporting the first of those suppositions can be found in the literature on GDP forecasting equations, see, for example, Muellbauer and Nunziata (2001), while evidence supporting the third can be found in Bowdler (2002). The second assumption is more controversial, for it is not clear that there is a systematic link between interest rate differentials and exchange rates. Still, Eichenbaum and Evans (1995) present econometric evidence indicating that the US dollar appreciates following major contractions of monetary policy, for example the Volcker deflation of the early 1980s. It is therefore surprising that empirical studies do not indicate at least some support for the idea that openness affects the slope of the Phillips curve.

We suggest two factors that may be responsible for past studies failing to detect a significant effect of openness on the slope of the Phillips curve: the measurement of the slope of the Phillips curve and the specification of the cross-sectional regressions intended to explain it.

Measuring the slope of the Phillips curve BMR measure the terms of the output-inflation tradeoff in a particular country through estimating the following regression over the period 1948-86:

$$y_t = const + \pi \Delta x_t + \lambda y_{t-1} + \gamma t \tag{1}$$

The log of real GDP, y_t, is regressed on a constant, its own lag, a time trend, and the change in the log of nominal GDP, Δx_t. The coefficient on the change in nominal demand, π, determines how much of a shock to nominal GDP in a particular year shows up in output, and is interpreted as a measure of the slope of the Phillips curve. An estimate of π close to unity indicates a very shallow Phillips curve in output-inflation space, while a value close to zero indicates a very steep Phillips curve. To verify this, note that if we define p as the log of the price level then we can use the fact that $x = p + y$ to rewrite (1) as follows:

$$\Delta p_t = \frac{1}{\pi}\left[(1-\pi)y_t + (\pi - \lambda)y_{t-1} - const - \gamma t\right] \tag{2}$$

Using equation (2) it is easy to show that the static elasticity of the inflation rate with respect to linearly detrended GDP (the output gap) is $(1-\lambda)/\pi$, which, *ceteris paribus*, is a decreasing function of π. This is consistent with BMR's claim that an estimate of π close to unity denotes a shallow Phillips curve in output-inflation space, while an estimate close to zero denotes a steep Phillips curve. A number of authors have criticised the BMR approach to measuring the slope of the output-inflation trade-off. Hutchison and Walsh (1998) note that equations (1) and (2) do not separately control for labour market shocks, exchange rate and raw material price fluctuations and movements in inflation expectations. Failure to identify those influences leads to their effects being incorporated into the estimate of π, such that the parameter no longer accurately measures the amount of inflation generated by a unit increase in the output gap. This point is recognised by Hutchison and Walsh:

> "..the estimated tradeoff, showing how nominal income changes are split between real output and price changes, will depend on the short-run output inflation tradeoff for a given expected rate of inflation (i.e. the slope of the Phillips curve) and the response of inflation expectations to changes in nominal demand (i.e. a shift in the Phillips curve)." [Hutchison and Walsh (1998), p. 712.]

In order to overcome the problems posed by measurement bias Hutchison and Walsh (1998) suggest augmenting models such as (2), with 'non-demand related explanatory factors determining inflation'. This is the approach followed by Bowdler (2002), who estimates inflation equations for 20 countries using quarterly data from the mid 1970s to the late 1990s, though varying slightly by country. Each equation is estimated separately and is obtained through testing down from the following baseline specification:

$$\begin{aligned}\Delta p_t \;=\;& \psi u_t + \sum_{m=1}^{5} \xi_m \Delta p_{t-m} + \sum_{j=1}^{6} \varsigma_j gap_{t-j} + \sum_{s=1}^{5} \vartheta_s \Delta ulc_{t-s} + \vartheta^* \left[ulc_{t-6} - p_{t-6}\right] + \\ & \sum_{r=1}^{5} \delta_r \Delta import_{t-r} + \delta^* \left[import_{t-6} - p_{t-6}\right] + \sum_{q=1}^{5} \alpha_q \Delta wpi_{t-q} + \\ & \alpha^* \left[wpi_{t-6} - p_{t-6}\right] + \sum_{w=1}^{5} \phi_w \Delta oil_{t-w} + \phi^* \left[oil_{t-6} - p_{t-6}\right] + \eta' D \qquad (3)\end{aligned}$$

where p is the price level, ulc is an index of unit labour costs, $import$ is an index of import prices, wpi is an index of wholesale prices, oil is the domestic currency price of oil (each of these variables being expressed in natural log form), D is a vector of dummy variables intended to remove the effects of outlying observations and ψ, ξ, ς, ϑ,

δ, α and ϕ are parameters and η is a vector of parameters.[8] The term *gap* measures the deviation of the natural log of GDP from its full employment potential, where potential GDP is modelled using an I(2) stochastic trend as in Aron and Muellbauer (2002), see Appendix A for full details of the procedure. In all applications the output gap term is a mean reverting process, ensuring that in the long-run output returns to its potential level, consistent with the notion that monetary policy cannot permanently raise GDP.

The model in (3) can be derived from a markup theory of the price level and is closely related to the specifications adopted by *inter alia* de Brouwer and Ericsson (1998), Aron and Muellbauer (2000) and Hendry (2001). It embeds a long-run solution for the price level of the form

$$p_t = \Psi_t + \vartheta' ulc_t + \delta' import_t + \alpha' wpi_t + \phi' oil_t$$

where Ψ_t is the component of u_t measuring the percentage markup of prices over a linearly homogeneous combination of input costs. Deviations from this long-run relation induce 'equilibrium corrections' in the price level that account for the local trends in inflation that complicate identification of the output-inflation tradeoff. Further, as the markup, Ψ_t, is potentially time-varying the empirical model will continue to account for drift in the inflation process following shifts in price-setting behaviour caused by changes in inflation expectations, see Bowdler (2002). The time-varying markup is fitted to the model as part of the local level term in (3), which is constructed from a random walk process using the STAMP package, see Koopman, Harvey, Doornik and Shephard (1995) and Bowdler (2002) for further details.

In order to measure the slope of the Phillips curve from these reduced form equations we invoke two key identifying assumptions. First, we assume that the equilibrium correction and local level terms jointly control for drift in the inflation rate that is unrelated to the output gap. This assumption is surely correct given that one can normally accept the hypothesis that the long-run pricing relation is linearly homogeneous, see the discussion in Bowdler (2002). Second, to deal with the fact that the output gap may raise inflation indirectly through its effect on factor markets we assume that the response of sectoral inflation rates to a 1% increase in the output gap is the same as the response of consumer price inflation, but that none of the inflation generated by the output gap arises via relative price changes.[9] Given these assumptions the full derivative of the inflation rate with respect to a 1% increase in the output gap, which constitutes our new measure of the slope of the Phillips curve (PC), can be calculated

[8] In a small number of cases limitations on data availability required some minor departures from this initial specification, see Bowdler (2002) for details.

[9] This is potentially a strong assumption. However, in Section 6 we show that relaxing it leaves the key empirical results unaffected, which suggests that its practical implications are quite limited.

as follows:

$$PC = \frac{\sum_{j=1}^{6} \varsigma_j}{\left[1 - \sum_{m=1}^{5} \xi_m - \sum_{s=1}^{5} \vartheta_s - \sum_{r=1}^{5} \delta_r - \sum_{q=1}^{5} \alpha_q - \sum_{w=1}^{5} \phi_w\right]} \tag{4}$$

We note three points at this stage. First, although PC is potentially a function of a large number of estimated parameters, such a possibility is more apparent than real, since the tested down inflation equations are very parsimonious, ensuring that (4) rarely incorporates more than four or five estimated coefficients.[10] Second, although we compute the elasticity of inflation with respect to the output gap out to infinity, very similar results are obtained through calculating the response out to, say, 6 quarters, since the sum of the coefficients in the denominator in (4) is typically quite small. Third, through excluding contemporaneous terms from the regression we constrain the inflationary impact of the output gap to be zero during the first three months. In practice this restriction is unlikely to be crucial since economic expansions typically affect price inflation with a lag, especially in quarterly data.

The correlations between PC[11] and the full sample and post-1973 BMR tradeoff measures are .33 and .25 respectively.[12] That these correlations are positive indicates some agreement between the two approaches as to which countries face relatively steep or relatively shallow Phillips curves. However, the associations are quite weak, a finding that we attribute to the measurement biases affecting the BMR methodology. For instance, the post-1973 BMR tradeoff measure is perversely signed for the UK, and generates a t-ratio of less than one for two other countries, Belgium and Norway, suggesting that the measurement bias is particularly severe in those cases. When the 20 countries in our sample are ranked according to the steepness of the Phillips curve using first PC and then the BMR parameter, the change in the rankings is 7 for the UK, 14 for Belgium and 15 for Norway, each of which exceeds the average shift of 6.6 places. Hence, the countries that account for the major differences between the two indices tend to be those for which the BMR method fares least well in identifying a conventional Phillips curve. This suggests that PC will be a more useful statistic in evaluating the impact of openness on the slope of the Phillips curve.

Specifying a cross-country regression A second potential reason for the absence of a correlation between openness and the slope of the Phillips curve is mis-specification

[10] Extensive dynamics are included in (3) because it represents the general specification used to identify the parsimonious model via general-to-specific modelling, see Bowdler (2002).

[11] The values taken by PC are available on request.

[12] These correlations refer to the negatives of the BMR tradeoff parameters, ensuring that increases in those indices describe a steepening of the Phillips curve, as is the case for increases in PC.

of the regression equations in Temple (2002). The theoretical discussion in Section 2 suggests that the slope of the Phillips curve is (potentially) determined as follows:

$$\begin{aligned} PC_i &= const + \beta_1 * OPEN_i + \beta_2 OPEN_i * SIZE_i + \beta_3 OPEN_i * EX_i + \beta_4 OPEN_i * TAY_i \\ &\quad + \gamma_1 INF_i + \gamma_2 INF_i^2 + \gamma_3 VOL_i + \gamma_4 VOL_i^2 \end{aligned} \tag{5}$$

where PC_i is the slope of the Phillips curve in country i, $OPEN_i$ measures the openness of country i, $SIZE_i$ is negatively related to the total GDP of country i, EX_i takes a relatively high value if country i sets monetary policy to stabilise the exchange rate, and TAY_i is the negative of the relative deviation of the time mean inflation rate of each country from the mean across countries.[13] INF_i is a measure of mean inflation and VOL_i is the measure of relative macroeconomic volatilities emphasised in the Lucas model (these terms are included as levels and squares to allow for non-linearities found to be important by BMR). Precise definitions of all of the variables are given in Section 4.

The cross-sectional models fitted by Temple are less general than (5) in that they implicitly assume $\beta_2 = \beta_3 = \beta_4 = 0$, thereby eliminating interaction terms from the regression. If those omitted terms are positively correlated with $OPEN$ and if $\beta_2, \beta_3, \beta_4 < 0$, as predicted by economic theory, then OLS estimation of (2) will yield a fitted value of β_1 that is biased towards zero. To investigate this possibility we undertake an empirical analysis of the relationship between openness and the slope of the Phillips curve that conditions on the full set of explanatory variables suggested by economic theory.

4 Data Collection

Unless otherwise stated, all data collected for use in this paper have been extracted from either the *International Financial Statistics* database maintained by the IMF or the *OECD National Accounts* available through *Datastream*. To measure $OPEN_i$ in (5) we compute the mean of the ratio of total import spending to nominal GDP in country i over the period for which an inflation equation was estimated for that country in Bowdler (2002).[14]

[13] The terms with which openness is interacted are defined in this way in order to ensure that β_2, β_3, and β_4 measure the extent to which the effect of openness in steepening the Phillips curve is turned off when, respectively, country i is too small to influence the international price of goods, fixes its exchange rate in order to avoid importing inflation, or has had such a low inflation rate in the past that the pass-through from exchange rate shocks to consumer prices is very limited.

[14] For Belgium, Greece, New Zealand and Sweden we averaged annual data over periods as close as possible to the quarterly periods studied in Bowdler (2002), i.e. always within one or two quarters.

In measuring $SIZE_i$ we first calculated the mean annual GDP of country i, measured in US dollars using 1995 prices and exchange rates, over the period for which an inflation equation was fitted for country i in Bowdler (2002). This figure was then subtracted from the mean of all such statistics for the sample of 20 countries, and the result divided by the cross-country mean. The resulting series varies negatively with a country's contribution to world GDP. When countries are too small to influence the international price of goods, $OPEN_i * SIZE_i$ will be relatively large. With $\beta_2 < 0$ there is then support for Romer's model in the sense that the effect of openness on the output-inflation tradeoff is 'turned off' in the case of a small country.

In constructing the variable EX we followed Campillo and Miron (1997) in classifying countries as following either fixed, semi-fixed or floating exchange rate regimes, and then assigned them a 0, 1 or 2 respectively (we use e to denote this indicator variable). Unfortunately, the Campillo-Miron classification refers to exchange rate regimes in 1974 and is therefore unsuitable for the present analysis. Instead, we obtained monthly data on the nominal effective exchange rate of country i over the period for which an inflation equation was estimated for that country in Bowdler (2002). Each series was then scaled by its mean and regressed on a constant and a time trend and the residual standard error was calculated. Figure A.1 in Appendix B shows these measures of exchange rate volatility in descending order. We then partitioned the sample into three groups corresponding to high, intermediate and low levels of exchange rate volatility, and used this classification to construct e (see Appendix B for the results).

The classification of countries across the three groups is broadly consistent with prior knowledge of the exchange rate regimes maintained by individual countries. For instance, the strict fixed exchange rate group comprises Germany and the smaller European countries that adhered most closely to the principles of the European Monetary System (EMS). The semi-fixed group mainly comprises the larger European countries whose currencies were less closely linked to the Deutsche Mark, e.g. Italy and the UK (both of whom eventually had to suspend membership of the EMS) and Spain and France (who remained a part of the EMS only through widening the target zones for their currencies), and also the Scandinavian countries, who opted for greater exchange rate flexibility following major macroeconomic shocks in the 1980s and 1990s (see Lindbeck (1997)). Lastly, the flexible exchange rate group mainly consists of non-European countries, since they have not participated in a scheme like the EMS. The exceptions to such rules are Greece (which appears in the floating group rather than the semi-fixed group), Canada (in the semi-fixed rather than the floating group), Norway (in the fixed rather than the semi-fixed group) and Austria (which appears in the fixed group even though, like Switzerland, it was not part of the EMS for the period that we

The exact sample periods are available on request.

consider). Such findings are consistent with the results in Calvo and Reinhart (2000), suggesting that these surprise outcomes are not the result of the particular methods that we employ. Rather, they reflect the fact that in practice exchange rate behaviour can deviate from a country's 'official' exchange rate policy. Clearly, it is the former concept that matters for determining the slope of the Phillips curve, and which should therefore be used to construct e. Still, in order to check that these ambiguous cases do not drive the central results reported in the next section, we excluded them from the analysis and found that this did not affect our main conclusions (results available on request).

The variable EX was then constructed from e as follows, in order to ensure that it has a zero mean and varies positively with the commitment to a fixed exchange rate:

$$EX_i = \frac{[e^* - e_i]}{e^*}$$

where e^* denotes the mean of e.

The interaction term $OPEN_i * EX_i$ will be relatively large for countries whose exchange rates tend to be fixed and which therefore should not import inflation to as great a degree over the course of the business cycle. Thus, if $\beta_3 < 0$, there is support for Lane's model of the relationship between openness and the slope of the Phillips curve, in that the effect of openness is 'turned off' in the case of a country that fixes its exchange rate.[15]

The interaction term TAY_i is obtained by subtracting the mean inflation rate for country i (for 1973q1 to the end of the period used to fit an inflation equation for country i) from the mean of such inflation rates for all 20 countries, and then dividing that figure by the mean. This ensures that $OPEN_i * TAY_i$ is high for countries that have experienced low inflation in the past. Then if $\beta_4 < 0$ it follows that the effect of openness on the slope of the Phillips curve is 'turned off' when low inflation conditions reduce the propensity of firms to pass exchange rate shocks through to consumer prices.

In order to test BMR's prediction we include INF_i and INF_i^2 in the cross-sectional regression. INF_i is the mean quarterly percentage inflation rate for country i from 1973q1 to the end of the sample period used for fitting an inflation equation for country i. Inflation rates are measured from the start of 1973 to allow for the fact that inflation may affect the frequency of contract negotiations with a lag.

The terms VOL_i and VOL_i^2 are included in the regression in order to test the predictions of the Lucas model. Both Lucas (1973) and BMR measure VOL_i as the standard deviation of the growth rate of annual nominal GDP, the argument being

[15]Strictly speaking, $\beta_3 < 0$ is also consistent with Romer's model, however, we can still distinguish between the two empirically if it is not possible to reject the hypothesis $\beta_2 = 0$.

that if the (unobservable) variance of preference shocks is the same across countries, then the slope of the Phillips curve should increase with the volatility of aggregate demand shocks. However, the standard deviation of nominal GDP growth will be affected by the volatility of real GDP growth, and is therefore a poor measure of nominal volatility. Instead, one should look at the volatility of unexpected movements in inflation. To calculate this we recursively estimate an AR(6) in the quarterly inflation rate for country i,[16] extract the one-step ahead forecast errors from 1985q1 onwards and define VOL_i as the standard deviation of that series (the residuals are calculated from 1985q1 onwards to allow sufficient observations for the initialization of the recursive estimation procedure).[17]

5 Empirical Results

The regression models in Table One examine open economy effects on the slope of the Phillips curve.[18] In order to make comparisons between past research and our own we use three measures of the slope of the Phillips curve: $-\pi_1$, the negative of the BMR tradeoff parameter calculated for 1948-86, $-\pi_2$, the negative of the BMR tradeoff parameter calculated for 1973-1986 and PC, the tradeoff measure described in Section 3.[19] As the regressand is always a derived variable, t-ratios are calculated using the heteroscedasticity consistent standard errors described in White (1980). The absolute values of these t-ratios are reported in parentheses in Table One. The results of a chi-square test for residual normality due to Doornik and Hansen (1994) are also quoted.

[16]This assumes that inflation expectations are formed adaptively. Ball (2000) justifies such an expectations process on the grounds that it implies small losses relative to the optimal inflation forecast and avoids the need for costly information collection.

[17]We multiply the calculated standard deviation by 100 to facilitate the estimation of the model, given that the original quarterly inflation rates were expressed as decimals.

[18]All regression estimates reported in this section and the next were obtained using the PcGIVE package of Hendry and Doornik (1999).

[19]We use the negatives of the BMR tradeoff measures in order to ensure that, like PC, the indices increase as the Phillips curve gets steeper in output-inflation space. Strictly speaking, one should use the reciprocal of BMR's tradeoff measure in making comparisons with PC. However, as the BMR parameter is actually negative for some countries, taking the reciprocal entails a non-monotonic transformation. To avoid this we simply multiply the BMR statistic by minus one.

Table One: Openness and the Output-Inflation Tradeoff				
Regression	(1)	(2)	(3)	(4)
Dependent Variable^A	$-\pi_1$	$-\pi_2$	$-\pi_2$	$-\pi_2$
Sample Size	20	20	20	18
$CONSTANT$	-.2755 (2.18)^B	-.5537 (4.51)	-.7354 (3.15)	-.7963 (3.50)
$OPEN$	-.1449 (.41)	.2891 (.87)	1.0159 (1.23)	1.2579 (1.61)
$OPEN * SIZE$				
$OPEN * EX$			-.4849 (1.02)	-.5861 (1.29)
$OPEN * TAY$				
Normality Test^C	1.65 (p = .44)	.48 (p = .79)	.69 (p = .71)	1.52 (p = .47)
R^2	.005	.02	.08	.17
Regression	(5)	(6)	(7)	(8)
Dependent Variable^A	PC	PC	PC	PC
Sample Size	20	20	20	20
$CONSTANT$	.1955 (5.28)	.0719 (1.87)	.0779 (2.14)	.0729 (1.49)
$OPEN$	-.1325 (1.31)	.3621 (2.58)	.3331 (2.64)	.3547 (1.67)
$OPEN * SIZE$				-.0088 (.10)
$OPEN * EX$		-.3300 (4.24)	-.2553 (3.67)	-.2597 (3.29)
$OPEN * TAY$			-.1641 (1.37)	-.1676 (1.35)
Normality Test^C	4.29 (p = .12)	.03 (p = .98)	1.34 (p = .51)	1.49 (p = .47)
R^2	.06	.60	.64	.64

A. π_1 is the BMR tradeoff parameter calculated for 1948-86. π_2 is the BMR tradeoff parameter calculated for 1973-1986. PC is the Bowdler tradeoff parameter.
B. Figures in parentheses are absolute t-ratios calculated using the heteroscedasticity consistent standard errors due to White (1980).
C. The normality test for the residuals is due to Doornik and Hansen (1994). The null hypothesis is that the residuals are normally distributed.

The model in column (1) confirms the main conclusion from past research, namely that openness is both incorrectly signed and insignificant in a bivariate regression for the terms of the output-inflation tradeoff that uses the full sample BMR tradeoff parameter. Further, that result is robust to restricting the sample from 42 countries to 20 countries.[20]

[20] The BMR and Bowdler samples actually have 19 countries in common, as New Zealand is included in the Bowdler study but not that by BMR. Here we are able to expand the sample to 20 through using a measure of the slope of the Phillips curve in New Zealand provided by Froyen and Waud (1995) using exactly the same data sources and econometric methods as BMR.

An obvious drawback in testing the Lane hypothesis using the 1948-86 BMR parameters is that they mainly refer to the Bretton Woods period, during which many countries maintained fixed exchange rates and thereby ensured that the impact of openness on the slope of the Phillips curve could not operate, at least not via the mechanism proposed by Lane. This problem can be overcome through replacing the full sample BMR tradeoff parameter with one estimated for the sub-period 1973-86. Column (2) shows that although the coefficient on openness takes the expected positive sign when $-\pi_2$ is the dependent variable, it is still insignificant. In column (3) we add the interaction between openness and the exchange rate regime. The slope coefficients are correctly signed in this model and the t-ratios are larger than in (2), but the effects are some way from achieving significance at the 5% level. A potential reason for this is that the BMR tradeoff parameter is subject to measurement bias, as argued in Section 3. To investigate this possibility we re-estimate the model after excluding the United Kingdom and Denmark from the sample. Those are the countries corresponding to the largest and smallest values of $-\pi_2$ respectively, and are therefore likely to be amongst the observations subject to the greatest amount of measurement bias. The results in column (4) indicate that the effect of openness is significant at the 15% level, suggesting that measurement bias may have been obscuring the link between openness and the slope of the Phillips curve in past studies.

In columns (5)-(8) we use PC as the dependent variable. Regressions (5) and (6) show that whilst there does not exist an unconditional relationship between openness and the slope of the Phillips curve, the expected positive effect does emerge after controlling for an interaction between openness and a country's exchange rate regime. Further, the inclusion of $OPEN * EX$ in the regression increases the R^2 statistic from .06 to .60, underlining its importance to the specification of the model. The importance of the interaction effect is due to the inclusion in the sample of several countries that have been in quasi-monetary union with Germany since the late 1970s. As the majority of monetary policy shocks in those countries have originated in Germany they have not induced exchange rate or import price adjustment, implying that the Phillips curves they have faced have not been as steep as their openness to trade would predict.[21]

This finding constitutes strong support for Lane's (1997) model of the output-inflation tradeoff, and reverses the result in Temple (2002). We conclude that the puzzle presented by Temple is jointly explained by measurement bias in the slope of the Phillips curve and not controlling for the effect of fixed exchange rate regimes.

[21]The coefficient estimates in column (6) indicate that fixing the exchange rate to the extent that countries such as Austria have done completely 'turns off' the effect of openness on the slope of the slope of the Phillips curve. However, we decline to draw such a strong conclusion here, on the grounds that actual parameter estimates may be quite poorly determined in very small cross-sectional samples.

Regression (7) in Table One indicates limited support for Taylor's model of moderated exchange rate pass-through. This contrasts with the strong support for the Taylor hypothesis reported by Choudhri and Hakura (2001), who found a positive cross-country correlation between mean inflation and an index of exchange rate pass-through. The failure of the $OPEN * TAY$ term to achieve significance is due to the inclusion of South Africa in the sample; 'dummying out' South Africa ensures that the regressor is significant at conventional levels (results not reported here). The dummy variable assigned to South Africa enters with a negative coefficient, indicating that its Phillips curve is not as steep as economic theory predicts. One potential reason for this is that price controls were used in South Africa over the period for which the slope of the Phillips curve was measured.

The final column in Table One shows that the point estimate on $OPEN * SIZE$ is correctly signed but highly insignificant. We therefore conclude that the theoretical model presented by Romer (1993) is less satisfactory than that due to Lane (1997).

Controlling for Closed Economy Effects We now analyse the effects of controlling for additional determinants of the output-inflation tradeoff. Table Two presents regressions comprising the variables emphasised in the Lucas and BMR models, as well as $OPEN$ and $OPEN * EX$.

Table Two: Openness, Inflation and the Output-Inflation Tradeoff

Regression	(1)	(2)	(3)
Dependent Variable	PC	PC	PC
Sample Size	20	20	20
$CONSTANT$	.1104 (6.65)	.0698 (1.85)	.0720 (1.94)
$OPEN$		.2980 (2.34)	.2719 (1.90)
$OPEN * EX$		-.2708 (3.83)	-.2521 (2.82)
INF^2	.0137 (3.62)	.0050 (1.10)	.0051 (1.05)
VOL^2			.0207 (.29)
Normality Test	.23 (p = .89)	1.49 (p = .48)	1.01 (p =.60)
R^2	.38	.63	.64

Notes: See notes B and C in Table One.

The model in column (1) regresses PC on a constant and the square of mean inflation (such a specification is preferred to one in which inflation enters linearly on grounds of best fit). The strong significance of the slope coefficient indicates that the basic spirit of the BMR study, namely that high inflation induces more frequent price-setting and a steepening of the Phillips curve, is robust to using a new measure of the slope of the Phillips curve.

This conclusion changes somewhat when we look at regression (2). The results indicate that whilst the open economy variables are significant in a regression that controls for the square of inflation, the inflation term itself is not. One possible reason for this is that we are relying on a small sample to identify effects from quite highly correlated variables. The third model in Table Two confirms BMR's finding that the Lucas model is of little help in explaining the slope of the Phillips curve.

In the final set of regressions presented in this section we follow Temple (2002) in conditioning on a series of variables describing labour market conditions. $WAGE$ $RIGIDITY$ is an index due to Grubb, Jackman and Layard (1983) that decreases with the speed of wage adjustment. The next two variables are taken from Bruno and Sachs (1985). $INDEXATION$ takes the value 0, 1 or 2 if wage indexation is, respectively, totally absent, partial or widespread, while $DURATION$ is also set to 0, 1 or 2, with higher values indicating relatively short price contracts. Hence, these two terms should enter the regression with a positive sign, while $WAGE\ RIGIDITY$ should enter with a negative sign. Observations on these variables are not available for all countries and so the sample size changes slightly across model specifications. For clarity, the exact sample size is quoted above each set of results.

Table Three: Models Incorporating Measures of Labour Market Inertia

Regression	(1)	(2)	(3)
Dependent Variable	PC	PC	PC
Sample Size	18	17	17
$CONSTANT$	.1258 (4.30)	.0938 (2.34)	.0930 (1.84)
$OPEN$	.2236 (2.09)	.3136 (1.92)	.2758 (2.29)
$OPEN * EX$	-.2609 (4.02)	-.3014 (4.18)	-.2837 (4.22)
$WAGE\ RIGIDITY$	-.0328 (5.09)		
$INDEXATION$		-.0075 (.55)	
$DURATION$			.0008 (.05)
Normality Test	.56 (p = .75)	1.95 (p = .38)	1.84 (p = .15)
R^2	.69	.58	.57

Notes: See notes B and C in Table One.

The results in column (1) indicate that greater wage rigidity leads to a reduction in the slope of the output-inflation tradeoff, as predicted by economic theory. In contrast, the other two variables are both insignificant, and the wage indexation term is incorrectly signed. Crucially from the point of view of this paper, open economy effects on the slope of the Phillips curve are robust to the inclusion of additional regressors.

6 Robustness Tests of the Empirical Results

In this section we examine the robustness of our core results. First, we report regressions obtained using the method of two-stage least squares. Second, we check that the results are not driven by outliers. Third, we present regressions in which the dependent variable is obtained using alternative identifying assumptions to those set out in Section 3.

Examining Regressor Exogeneity The estimated coefficients in Section 5 may capture the endogenous responses of the conditioning variables to the slope of the output-inflation tradeoff. For example, suppose a country faces a relatively flat short-run Phillips curve, e.g. due to high levels of wage rigidity. For a given sequence of aggregate demand shocks this country will generate a relatively low variance inflation process and therefore a low variance detrended exchange rate, suggesting that the $OPEN * EX$ term could be negatively signed even if the Lane theory is irrelevant to the determination of the output-inflation tradeoff.

To deal with potential endogeneity problems we consider regressions estimated via two-stage least squares (2SLS). The model that we concentrate on is that in which PC is regressed on a constant, $OPEN$ and $OPEN * EX$. We treat $OPEN$ as endogenous in column (1), but exogenous in column (2). In contrast, variation in $OPEN * EX$ is treated as endogenous in both cases. We use lagged openness and its square as instruments in the first stage regressions in column (1), and in column (2) $OPEN$ is available as a further instrument.[22] The absolute t-ratios given in parentheses are based on the corrected standard errors computed by the *PcGIVE* package, see Doornik and Hendry (1999). The R^2 statistic is not uniquely defined for 2SLS estimates, so here we report the regression standard error as a measure of the fit of each specification.

[22]Lagged openness is calculated as the mean of openness from the start of 1970 to the quarterly time period immediately before that in which the estimation of an inflation equation for a particular country begins. In some cases we averaged annual observations on openness. The exact sample periods are available on request.

Table Four: Regressions for the Tradeoff Parameter Estimated via 2SLS		
Regression	(1)	(2)
Dependent Variable	PC	PC
Sample Size	20	20
$CONSTANT$	-.0389 (.39)^{A}	-.0048 (.07)
$OPEN$	.8002 (2.09)	.6687 (2.47)
$OPEN * EX$	-.6010 (2.70)	-.5345 (3.18)
Normality Test	.04 (p = .98)	.05 (p = .98)
Standard Error	6.51%	5.78%
A. The t-ratios quoted here are based on standard errors corrected for 2SLS estimation.		

The results in columns (1) and (2) indicate that our core findings are robust to estimation of the model by 2SLS - the significance of openness and the interaction between openness and the exchange rate regime does not appear to be driven by reverse causation bias.

Examining the Role of Outliers When testing macroeconomic theories using a sample of just 20 countries it is important to check that the core results are not driven by outliers. We therefore consider the plot of PC against $OPEN$ after having first regressed each variable on $OPEN * EX$ and a constant (we denote these variables PC' and $OPEN'$), and then the plot of PC against $OPEN * EX$ after having first regressed each variable on $OPEN$ and a constant (we denote these variables PC'' and $OPEN * EX''$). These plots, together with lines of best fit, are presented in Figure One.

Figure One - see end of document.

An inspection of the plots in Figure One suggests that the relationship between openness, the exchange rate regime and the output-inflation tradeoff is due to the average information in the sample. We therefore interpret our core findings as a central feature of macroeconomic adjustment even though they are derived from a small sample.

Using Alternative Tradeoff Measures As a final robustness check we re-estimate our basic model using three alternative measures of the slope of the Phillips curve. The first is obtained through applying the formula in (4) to inflation equations estimated using data spanning just the first halves of the periods studied in Bowdler (2002) and is denoted $PC1$. The second measure ($PC2$) is obtained through deleting the local level terms from the inflation equations reported in Bowdler (2002), replacing them with the level, square, cube and fourth power of a time trend and then applying the OLS estimator.[23] The formula in (4) is then used to obtain the new measure of the slope of the Phillips curve.

Finally, we obtain $PC3$ through relaxing our earlier assumption that the output gap impacts input price inflation rates in the same way that it impacts consumer price inflation. Instead, the strength of these indirect effects is freely estimated from the data. To do this we first estimate new reduced form inflation equations through testing down from the following specification:

$$\begin{aligned}\Delta p_t \quad = \quad & \psi u_t + \sum_{m=1}^{5} \xi_m \Delta p_{t-m} + \sum_{j=1}^{6} \varsigma_j gap_{t-j} + \sum_{s=1}^{6} \vartheta_s \left[ulc_{t-s} - p_{t-s}\right] + \\ & \sum_{r=1}^{6} \delta_r \left[import_{t-r} - p_{t-r}\right] + \sum_{q=1}^{6} \alpha_q \left[wpi_{t-q} - p_{t-q}\right] + \\ & \sum_{w=1}^{6} \phi_w \left[oil_{t-w} - p_{t-w}\right] + \eta' D \end{aligned} \tag{6}$$

This yields inflation equations that condition on equilibrium correction terms, the local level term and lags in inflation and the output gap. Crucially, no terms in sectoral inflation rates are included in these models.[24] Given our first identifying assumption, namely that the local level and equilibrium correction terms account for only the underlying trend in inflation, the amount of inflation generated by a 1% increase in the output gap is calculated as follows:

$$\frac{\sum_{j=1}^{6} \varsigma_j}{\left[1 - \sum_{m=1}^{5} \xi_m\right]} \tag{7}$$

As $PC3$ is calculated from inflation equations that do not hold sectoral inflation rates constant, the amount of inflation generated through the output gap pushing up prices in labour and raw materials markets is captured by the static partial elasticity of the inflation rate with respect to the output gap, i.e. the quantity in (7).

[23]The time trend used to create these terms was divided by 100 to facilitate the estimation of the model.

[24]In obtaining the inflation equations underpinning $PC3$ we had to make a change to the standard general-to-specific modelling strategy. Consider a case in which a term like $\vartheta \Delta ulc$ enters one of the inflation equations estimated by Bowdler (2002). Such a term can also enter the reduced form obtained from (9), and would appear as $\vartheta \Delta \left[ulc - p\right] + \vartheta \Delta p$. In order to avoid obtaining reduced forms of this sort, in cases in which consecutive lags of a particular equilibrium correction term entered an estimate of (9) with opposite signs (suggesting the reparameterisation $\Delta \left[ulc - p\right]_{t-s}$) we deleted them from the model.

In Table Five we report regressions of each of the new measures of the slope of the Phillips curve on a constant, openness and the interaction of openness and the exchange rate regime. We measure these regressors over exactly the same periods as in Sections 4 and 5, even in the case of $PC1$, the version of the dependent variable based on the sub-sample inflation equations. This is unlikely to have a major impact on the results because the explanatory variables change relatively little over time.

Table Five: Regressions Using $PC1$, $PC2$ and $PC3$			
Regression	(1)	(2)	(3)
Dependent Variable	$PC1$	$PC2$	$PC3$
Sample Size	20	20	20
$CONSTANT$	-.0194 (.25)	.0305 (.34)	.0796 (2.61)
$OPEN$	.9261 (2.68)	.5939 (1.46)	.2033 (2.17)
$OPEN * EX$	-.7155 (3.58)	-.4729 (1.74)	-.2018 (5.25)
Normality Test	5.07 (p = .08)	9.46 (p = .01)	1.08 (p = .58)
R^2	.55	.31	.47
Notes: See notes B and C in Table One.			

The results in column (1) indicate that our key findings are robust to measuring the terms of the output-inflation tradeoff using data up to the mid/late 1980s. The coefficient estimates are substantially higher than those reported in Table One where we used PC as the dependent variable, something that can be traced to the Phillips curve in Greece being more than twice as steep when measured by $PC1$ rather than PC. Excluding Greece from the analysis yields a set of coefficient estimates much closer to those recorded in Table One (results not reported here). The second column in Table Five indicates that the effect of openness is somewhat less significant when the slope of the Phillips curve is measured using time series equations fitted by least squares. This loss of significance is mainly due to an increase in the standard errors attached to the coefficients, however, and the magnitudes of the estimated effects actually increase compared to those in Table One. It seems that the slope of the Phillips curve is measured with less precision when we control for unobserved shifts in the inflation rate via deterministic terms rather than a local level.

Finally, the results in the third column indicate that our main finding is robust to relaxing our assumption about the way in which the output gap raises inflation through its impact on goods market prices. The main reason for this is that the coefficients on lagged sectoral inflation rates in the equations estimated by Bowdler (2002) are typically quite small, perhaps in the range 0 to 0.1. As a result, it is the coefficients multiplying the output gap that do most of the work in determining both PC and $PC3$.

7 Summary

The principal finding described in this paper is that increased openness to trade increases the slope of a country's short-run output-inflation tradeoff, or Phillips curve, provided that the exchange rate of that country is free to adjust to shifts in monetary policy. Such a condition is crucial, for it ensures that fluctuations in economic activity can induce the changes in import prices necessary to accelerate inflation adjustment. Such a result is consistent with the model of the output-inflation tradeoff in Lane (1997).

The puzzle presented by Temple (2002), namely that openness exerts no impact on the slope of the Phillips curve, disappears when one uses better measures of the slope of the Phillips curve and conditions on the exchange rate regime.

We found no evidence to support Romer's theory of the Phillips curve. Such a model is only potentially relevant to the United States, the world's largest economy, and actually seems to have little practical significance for that country, since the United States has one of the flattest Phillips curves of the 20 countries that we studied. On the other hand, we found partial support for Taylor's (2000) model of moderated exchange rate pass-through.

Appendix A: Measuring Trend GDP

The time series equations reported in Bowdler (2002) are used as a basis for measuring the slope of the Phillips curve in Section 3 above. Those equations make use of an output gap series based on a measure of trend output obtained using a stochastic trend technique closely related to that employed in Aron and Muellbauer (2002) and available as part of the STAMP package of Koopman et al (1995). This appendix provides brief notes on that technique. The total level of GDP, y_t, is modelled as the sum of a smooth trend (χ_t), a trigonometric function ($\varkappa_t$) and an error term (ε_t), i.e. we have

$$y_t = c + \upsilon\chi_t + \varrho\varkappa_t + \varepsilon_t, \quad \varepsilon_t \sim NID\ (0, \upsilon_\varepsilon^2)$$

$$\chi_t = \chi_{t-1} + \nu_{t-1} + \iota_t, \quad \iota_t \sim NID\ (0, \upsilon_\iota^2)$$

$$\nu_t = \nu_{t-1} + \varpi_t, \quad \varpi_t \sim NID\ (0, \upsilon_\varpi^2)$$

$$\begin{bmatrix} \varkappa_t \\ \varkappa_t^* \end{bmatrix} = p_\varkappa \begin{bmatrix} \cos\Gamma & \sin\Gamma \\ -\sin\Gamma & \cos\Gamma \end{bmatrix} \begin{bmatrix} \varkappa_{t-1} \\ \varkappa_{t-1}^* \end{bmatrix} + \begin{bmatrix} \kappa_t \\ \kappa_t^* \end{bmatrix}$$

where c is a constant, $p_\varkappa$, $0 < p_\varkappa \leq 1$, is a damping factor, Γ is the frequency (in radians) of the cyclical term, and κ_t and κ_t^* are two mutually uncorrelated NID disturbances with zero mean and common variance υ_κ^2. The estimation of the model proceeds in two steps. A maximum likelihood technique is used to compute estimates of the unknown variances and then the Kalman filter is passed through the data in order to give the estimated coefficients.

Trend GDP is defined as $c + \upsilon\chi_t$, and the output gap is measured as $y_t - c - \upsilon\chi_t$. This measure of the output gap is to be preferred to the Hodrick-Prescott measure, for it does not rely on any arbitrary calibration of the variance of the trend term. Further, the problem of excessive variation in the trend towards the end of the sample that is known to affect the Hodrick-Prescott method is less severe in the present case due to the presence of the trigonometric term, which captures cyclical variation in the data and therefore restricts movements in the trend.

Appendix B: Notes on the Construction of the 'EX' Dummy

Figure A.1 presents measures of the volatility of linearly detrended nominal effective exchange rate data for each of the 20 countries that we study (see the text for notes on the computation of these statistics). The abbreviations used in Figure A.1 are as follows: AUS=Australia, AU=Austria, BE=Belgium, CA=Canada, DE=Denmark, FI=Finland, FR=France, GE=Germany, GR=Greece, IT=Italy, JA=Japan, NE=Netherlands, NZ=New Zealand, NO=Norway, SA=South Africa, SP=Spain, SW=Sweden, SWI=Switzerland, UK=United Kingdom, US=United States.

Figure A.1 - see end of document.

We choose to divide the sample into three sub-groups, each corresponding to a different level of exchange rate volatility. These are indicated by the solid dividing lines in Figure A.1. To be sure, the exchange rate regime indicator, e, is set to 0 for Austria, Belgium, Germany, the Netherlands and Norway (the fixed exchange rate group), to 1 for Canada, Denmark, Finland, France, Italy, Spain, Sweden, Switzerland and the UK (the semi-fixed exchange rate group), and to 2 for Australia, Greece, Japan, New Zealand, South Africa and the United States (the flexible exchange rate group).

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Figure One: Conditional Effects on the Slope of the Phillips Curve

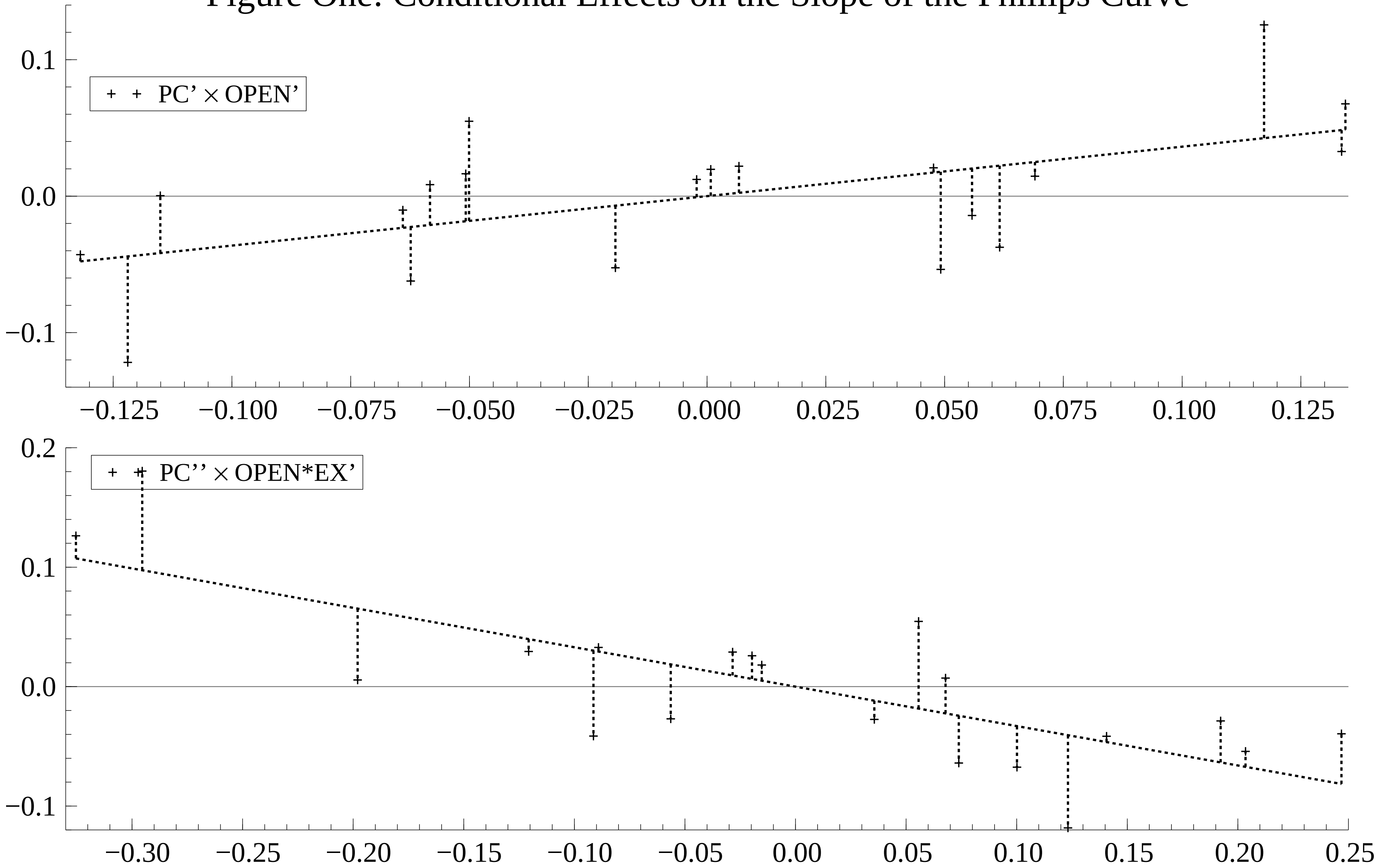

Figure A.1: Detrended Exchange Rate Volatilities