Openness and the Output-Inflation Tradeoff*

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November 4, 2003

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 indicate some support for the standard theoretical prediction, but it is confined to those countries that have maintained floating exchange rate regimes.

Keywords: Openness, Inflation, Phillips curve.
JEL Classification: E31, E32, F41.

1 Introduction

This paper tests the hypothesis that the slope of the short-run Phillips curve, defined as the amount of inflation generated by a unit increase in output relative to trend, varies positively with openness to trade. A series of cross-sectional regressions are presented, 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 results indicate that international differences in openness to trade exert a positive effect on the slope of the Phillips curve (or output-inflation tradeoff) provided that the countries concerned maintain flexible exchange rate regimes.

This conclusion is at odds with the evidence presented in Temple (2002), which indicates that openness does not exert a systematic effect on the slope of the Phillips curve. We suggest two

*This research was completed during the first year of a British Academy post-doctoral fellowship. I am very grateful to the Academy for financial support. I would like to thank John Muellbauer for the excellent advice that he has provided whilst I was undertaking this research. Special thanks are also due to Jon Temple for many helpful comments and suggestions. Further, I have benefited from comments on a previous draft of the paper from Rui Fernandes, Roland Meeks, Heino Bohn Nielsen, Adrian Pagan, Phurichai Rungcharoenkitkul, Kamakshya Trivedi and seminar participants in Oxford and at the 2003 Money, Macro and Finance conference in Cambridge. Any remaining flaws are my responsibility.

reasons for the differences between past results and our own. First, previous research has been based upon a measure of the slope of the Phillips curve due to Ball, Mankiw and Romer (1988), hereafter BMR. 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 may therefore provide biased measures of the slope of the Phillips curve. In order to deal with this problem, we replace the BMR parameter with an alternative measure of the slope of the Phillips curve derived from a three equation model, which we estimate for 19 countries. 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 exchange rate regime type. We incorporate this interaction term into the cross-sectional regression analysis and find that it is very important in identifying the strength of the relationship between openness and the slope of the Phillips curve.

The robustness of the evidence linking openness, the exchange rate regime and the slope of the Phillips curve is evaluated in some detail in the second half of the paper. The strength of the association turns out to be somewhat sensitive to changes in the set of countries used in the testing procedure. Specifically, the relationship is confined to those OECD countries that have followed flexible exchange rate policies. As this group comprises just one quarter of the full set of countries in the sample, excluding a small number of countries from the analysis can substantially weaken the main relationships identified using the full sample.

The remainder of the paper expands on these points and has the following structure. Section 2 discusses the underlying economic theory. Section 3 reviews some empirical tests of the theoretical predictions, particularly focusing on issues relating to measurement and model specification. Section 4 reports the new empirical results that we obtain, and Section 5 rounds off with a summary.

2 Economic theory

The theoretical basis for the hypothesis that the slope of the Phillips curve is related to openness derives from the contributions of Romer (1993) and Lane (1997), and is described in non-technical terms in Temple (2002). In Appendix A we sketch out the details of the Romer model. The focus is an open economy in which the average price of domestically produced goods is sticky and therefore adjusts only gradually in response to supply and demand shocks (this could be due to the effect of overlapping contracts or heterogeneous costs of price adjustment). In such a model, the policy authority can adjust the money supply in order to manipulate output over the short-term, i.e. expansions of the money supply raise both output and prices in the short-run, such that there is a positively sloped output-inflation tradeoff. Only in the long-run, when all nominal variables have been set to their equilibrium values, does the Phillips curve take a vertical form in output-inflation space.

Romer argues that monetary policy expansions are associated with depreciation of the nominal exchange rate (the basis for this claim will be discussed shortly). This affects the slope of the Phillips curve over the short-term through two separate mechanisms. First, when inflation is measured in terms of a consumer price index, the effect of the depreciation will be to add to the amount of inflation associated with a particular change in the money supply. Second, if wages are partially indexed to a consumer price index, or if foreign goods are used as intermediate inputs in domestic production, the output gain to a given monetary expansion will be reduced. Clearly, both effects will be stronger in more open economies, because the share of imported goods and services in both the consumer price index and the producer price index is larger in such cases. This implies that more open economies will face steeper Phillips curves in output-inflation space (see Appendix A for a derivation of this result within a simple open economy model).[1]

The pivotal assumption in Romer's model is that the monetary policy expansions that drive the output gap also cause nominal exchange rate depreciations (and also real exchange rate depreciations, given short-run domestic price stickiness). In the model analysed by Romer, this follows from the fact that each country is large enough to influence the international price of goods through its own supply and demand decisions. Specifically, if domestic and overseas output are imperfect substitutes in domestic consumption, then a monetary expansion will raise the desired quantity of imports. In order that these imports be acquired, extra domestic output must be supplied to the world market. The price of domestic output relative to foreign output must then fall in order to clear international markets, i.e. there is a depreciation of the real exchange rate. As this feeds into higher import prices in the country in which the policy shock occurred, a positive correlation between openness and the slope of the Phillips curve is expected.

A difficulty with this argument lies in its first step: most countries are not large enough to influence international relative prices simply by expanding domestic output.[2] In order to overcome this problem, Lane (1997) modifies Romer's model so that exchange rates are determined by arbitrage in foreign exchange markets, as in the Mundell-Fleming model. As is well known, increases in the money supply generate exchange rate depreciations in this model, while decreases in the money supply generate exchange rate appreciations. Given that prices are sticky, these nominal exchange rate dynamics imply that the real exchange rate will depreciate when the output gap is positive. Embedding these responses in the Romer model ensures that a country characterised by greater trade openness will face a steeper Phillips curve irrespective of whether or not it is large enough to influence international prices through its supply decisions.

The Lane analysis points to one important caveat concerning the relationship between trade openness and the slope of the Phillips curve: If a country maintains a fixed exchange rate regime, i.e. it does not set monetary policy independently of major trading partners, then expansions of monetary policy (and hence the output gap) will not be associated with an acceleration of the import price index, and we would not expect to observe a relationship between trade openness and the slope of the Phillips curve. In the remainder of the paper, we take this into account through testing the following conditional prediction:

The slope of the Phillips curve will vary positively with openness to trade, but the

[1] It should be noted that 'openness' in this context refers to import penetration, as opposed to any alternative definition based upon the structure of trade barriers, or the mobility of financial capital.

[2] 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 the ability to influence world prices.

strength of this association will decrease as a country's monetary policy authority increases its commitment to fixing the exchange rate.

A final point that should be noted in relation to the Romer-Lane hypothesis is that the argument assumes that movements along the short-run Phillips curve are predominantly driven by monetary shocks, rather than fiscal policy. This may be a reasonable approximation to the reality of the past 25 years, for during that time fiscal policy has been subordinated to a largely microeconomic role, and monetary policy has been used as the main instrument of macroeconomic control. Nevertheless, the assumption should be borne in mind when considering the results reported in this paper.

3 Testing the Romer-Lane hypothesis

The impact of openness to trade on the slope of the Phillips curve has been tested by Temple (2002), using cross-country regression analysis applied to a sample of 42 countries. The results indicate that openness (measured as the share of imports in GDP) exerts an insignificant effect on the slope of the Phillips curve, and that the estimated relationship is of the opposite sign to that predicted by economic theory. The robustness of this finding is confirmed using a least trimmed squares estimator, through changing the time period over which the slope of the Phillips curve is measured and through augmenting the regressions with further control variables.

The reasons for the lack of empirical support for the Romer-Lane prediction are not clear. Recall that the key steps in the argument are that monetary policy expansions should both raise output and depreciate the exchange rate (and vice versa), and that exchange rate driven fluctuations in import prices be passed through to consumer prices. Empirical evidence suggesting that monetary policy affects output can be found in the literature on GDP forecasting equations, see, for example, Muellbauer and Nunziata (2001), while the empirical relevance of the third part of the mechanism is demonstrated in Hendry (2001). It is much less clear that there exists a systematic link between the stance of monetary policy and the value of the exchange rate, see Obstfeld and Rogoff (1996, p. 621-622). Nevertheless, 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. This suggests that following major policy interventions the pattern of macroeconomic adjustment necessary for a correlation between openness and the slope of the Phillips curve does apply, and it is therefore surprising that empirical studies do not indicate at least some support for the idea.

In this section we suggest two factors that may account for the lack of a correlation between openness and the slope of the Phillips curve in past studies: 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

The BMR estimate of the slope of the Phillips curve in a particular country is obtained by fitting the following regression using annual data for 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:

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

In equation (2) the increase in inflation during a year in which a unit shock to output occurs (after controlling for the linear trend in both variables) is $(1-\pi)/\pi$. As this magnitude is decreasing in π for $\pi > 0$, it follows 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 this approach to measuring the slope of the Phillips curve. Akerlof, Rose and Yellen (1988) shows that OLS estimates of equation (1) are subject to a simultaneity bias. Hutchison and Walsh (1998) argue that the omission of wage and raw material price effects from (1) implies that the estimated π coefficient may be distorted by supply-side shocks that affect both output and inflation. For example, an oil price hike may be expected to raise inflation and decrease output, such that estimates of π are biased towards zero. In practice, π is estimated to be negative rather than positive for approximately one quarter of the countries studied by BMR. Such a finding is inconsistent with standard formulations of the Phillips curve, in which output and inflation are positively associated, and suggests that the BMR index is subject to important measurement biases that may obscure the relationship between openness to trade and the slope of the Phillips curve.

In this paper we construct new measures of the slope of the Phillips curve for 19 countries.[3] The basic idea is to measure the amount of inflation associated with a unit shock to output *after controlling for the non-demand related movements in those two variables.* The starting point for the analysis is the following set of equations:

[3] This is the set of developed countries amongst the 42 country sample studied by Temple (2002), excluding Hong Kong, Ireland and Portugal, for which we could not obtain sufficient time series data. The sample that we use does not extend to developing countries because past studies suggest that the inflation process in developing countries cannot be explained using linear models based on the information set that we utilise in our analysis, see, for example, Aron and Muellbauer (2000).

$$
\begin{aligned}
\Delta p_t &= const + \sum_{m=1}^{4} \xi_m \Delta p_{t-m} + \sum_{j=1}^{4} \varsigma_j gap_{t-j} + \sum_{s=1}^{4} \vartheta_s \left[ulc_{t-s} - p_{t-s}\right] \\
&\quad + \sum_{w=1}^{4} \phi_w \left[usoil_{t-w} - p_{t-w}\right] + \beta_1' t + \beta_2' t^2 + \beta_3' t^3 \qquad (3) \\
gap_t &= const + \sum_{j=1}^{4} \varsigma_j' gap_{t-j} \qquad (4) \\
[ulc - p]_t &= const + \sum_{s=1}^{4} \vartheta_s'' \left[ulc_{t-s} - p_{t-s}\right] + \sum_{j=1}^{4} \varsigma_j'' gap_{t-j} \qquad (5) \\
&\quad + \sum_{w=1}^{4} \phi_w'' \left[usoil_{t-w} - p_{t-w}\right] + \beta_1'' t + \beta_2'' t^2 + \beta_3'' t^3
\end{aligned}
$$

The variables in equations (3)-(5) are in natural log form, and observed at the quarterly frequency. The variable definitions are as follows: p denotes the price level, gap is the deviation of output from trend, ulc is unit labour costs, $usoil$ is the US$ price of oil and t is a time trend. The procedure that we follow in deriving a measure of the slope of the Phillips curve from this set of equations is as follows:

1. Equations (3)-(5) are estimated separately by OLS using quarterly data running from the late 1970s to the late 1990s[4] (the exact sample periods differ slightly across countries, see Appendix B for further details and for information concerning the measurement of the variables in (3)-(5)).

2. Each equation is reduced to a parsimonious form using the *PcGETS* algorithm developed by Hendry and Krolzig (2001). This is a computer programme that carries out automated reduction of highly parameterised time series equations to their minimum dimensions, through repeated estimation of the equation by OLS. The criteria used in implementing this general-to-specific modelling strategy include the individual and joint significance of the variables, and the outcomes of residual diagnostic tests and tests for parameter stability.

   The programme offers the choice between a 'liberal' strategy, which minimises non-selection of variables, and a 'conservative' strategy, which minimises non-deletion. The liberal strategy was adopted in this exercise. An outlier correction procedure (available as part of the programme) was also used. This assigns a dummy variable to extreme observations so that they cannot distort the results, see Hendry and Krolzig (2001) for details.

3. The tested down versions of (3)-(5) were then re-estimated as a three equation system using the Full Information Maximum Likelihood (FIML) estimator

[4] The sample periods begin after the dissolution of the Bretton Woods fixed exchange rate regime in 1973 and end before the introduction of the single European currency in January 1999.

available in the *PcGIVE10* package of Hendry and Doornik (2001). This systems method increases the efficiency of the estimation.[5]

4. The three equation systems were then used to compute the response of inflation to a unit increase in gap_t at horizons $t+1$, $t+2$, $t+3$ and $t+4$ using the impulse response command in *PcGIVE10.* The sum of those four quarterly responses defines the new measure of the slope of the Phillips curve, and is denoted PC.

We note two features of the procedure used in testing down from the general model in (3)-(5) to the specific model used for calculating the impulse responses. First, as it uses information on equation stability and the properties of the residuals, the final specification that it delivers provides a reliable basis for measuring the dynamic impact of the output gap on inflation. Alternative models that yield serially correlated errors or unstable coefficients are clearly less reliable - see Hendry and Krolzig (2003) for further details on the properties of *PcGETS*. Second, as the procedure is automated, the possibility that arbitrary reductions of equations (3)-(5) deliver estimated Phillips curve parameters that correlate with openness purely by chance is avoided.[6]

The impulse response analysis used to construct the PC index measures two types of effect. First, in equation (3), the impact of the output gap on inflation is evaluated holding constant all fluctuations in real unit labour costs and US$ oil prices relative to domestic prices.[7] As these two conditioning variables are likely to capture the effects of major supply shocks arising in labour markets and commodity markets, the partial derivative calculated from (3) is less likely to be affected by omitted variable bias than is the BMR parameter. Second, equation (5) is used to measure the amount of inflation arising indirectly through the output gap feeding into real unit labour costs, and real unit labour costs then entering the equation for inflation. Through combining these two effects, PC measures the full derivative of inflation with respect to the output gap.

The US$ price of oil, *usoil*, is a non-modelled variable in (3)-(5), which means that the output gap does not affect inflation through first raising the US$ price of oil. This is a reasonable assumption because any single country normally represents a small share of the global oil market, and is therefore unlikely to be able to influence US$ oil prices. An obvious exception to this rule is the United States itself, of course. In order to cast some light on the size of demand-side contributions to real oil prices in the United States, we ran an OLS regression of real oil prices on four lags in real oil prices and four lags in the output gap. None of the output gap coefficients turned out to be significant, even when we tested the model down to a parsimonious form, and we therefore continue to treat oil prices as exogenous in the determination of the output-inflation tradeoff.

[5] The systems estimator cannot be applied to the general-to-specific modelling because the *PcGETS* software does not yet allow for such an approach.

[6] All of the computer output corresponding to this section of the paper is available on request.

[7] The use of input costs relative to domestic prices (as opposed to the first differences of input prices) as a means of controlling for supply-side influences on inflation follows the 'error correction' approach to modelling inflation fluctuations, see, for example, de Brouwer and Ericsson (1998).

The most general forms of the inflation and real unit labour cost equations contain a cubic in time. These deterministic terms are intended to capture any non-demand related influences on those two variables that are not controlled for by the other explanatory variables, e.g. changes in product market structure that affect inflation via the price-cost markup, or changes in trade union power that affect real unit labour costs. Finally, note that all contemporaneous terms are excluded from (3)-(5) in order to ensure that the regressor set is pre-determined and that simultaneity biases do not affect the estimation. One disadvantage of this approach is that the Phillips curve is constrained to be completely flat during the quarter in which a shock occurs, and can only exhibit a positive slope in the four subsequent quarters. In contrast, the BMR method measures the response of inflation to the output gap over four quarters, including that quarter in which a shock occurs.

The PC parameters for the 19 countries that we study in this paper are listed in Appendix C alongside the versions of the BMR statistic calculated using annual data for 1973-86. In contrast to the BMR index, the PC parameter is non-negative, suggesting that the multivariate techniques used to control for the influence of supply-side shocks have been effective. The Phillips curve parameters range from 0 to 1.309, and the mean parameter value across the 19 countries is 0.447, indicating that a 1% increase in GDP relative to trend adds, on average, approximately 0.5% to the annual inflation rate in the first year. As a simple check on the robustness of the results, we also calculated total inflation responses to a unit shock to the output gap in period t over the horizon $t+1$, ... , $t+8$. The correlation between the two indices was .845, indicating that the pattern of international differences in the slope of the Phillips curve is not very sensitive to changes in the time horizon over which inflation responses are calculated.

In six cases the slope of the Phillips curve is found to be zero during the first four quarters following a shift in the output gap. In three of these instances, Italy, the Netherlands and Sweden, this is due to the output gap raising inflation indirectly via its effect on real unit labour costs, and the total lag in that effect exceeding four quarters. In the other three cases, Austria, Denmark and Spain, there appears to be no Phillips curve relation at any horizon. One reason for this may be that the output gap affects inflation with a lag of more than four quarters in those three countries. Allowing each of these six countries to have a positively sloped Phillips curve through adding gap_{t-1} to the inflation equation leads to a set of PC parameters that has a 99.65% correlation with the original set. This suggests that the horizontal Phillips curves that we identify are not an artefact of the model selection procedure.[8]

The correlation between the PC index and the BMR parameter measured over 1973-86 is -.341, which indicates some agreement between the two approaches as to which countries face relatively steep Phillips curve (recall that the BMR index decreases with the slope of the Phillips curve, so a negative correlation is to be expected if the two indices are in agreement).[9] However,

[8] The absence of a Phillips curve effect in many of these countries is confirmed in a more detailed study of OECD inflation featuring higher order dynamics and unobserved components in the estimation process, see Bowdler (2003).

[9] The correlation statistic that has been calculated understates the amount of agreement across the two approaches because (i) the sample periods are not identical; (ii) PC_i is comparable with $[\frac{1-\pi}{\pi}]_i$, not π_i (we cannot calculate $[\frac{1-\pi}{\pi}]_i$, however, because π_i is estimated to be less than zero in some cases, such that the required transformation is non-monotonic and therefore would not produce meaningful results).

the correlation coefficient is still very far from minus one, suggesting that a test of the Romer-Lane hypothesis based upon the PC index may yield different conclusions to those reached by Temple (2002) using the BMR parameter.

Specifying a cross-country regression

A second potential reason for the absence of a correlation between openness and the slope of the Phillips curve in Temple (2002) is that the cross-sectional regression models employed are not general enough. The theoretical discussion in section 2 suggests that the slope of the Phillips curve is (potentially) determined as follows:

$$PC_i = const + \beta_1 * OPEN_i + \beta_2 * (OPEN_i * EX_i) \tag{6}$$

where PC_i is the slope of the Phillips curve in country i, $OPEN_i$ measures the openness of country i and EX_i measures the extent to which monetary policy in country i is set to stabilise the exchange rate, i.e. EX_i takes a relatively large value if country i maintains a fixed exchange rate regime. If the predictions of the Lane model are correct then β_2 will be estimated to be negative, and the magnitude of β_2 will measure the extent to which the effect of openness in steepening the Phillips curve is 'turned off' when country i fixes its exchange rate. The cross-sectional models fitted by Temple implicitly assume $\beta_2 = 0$, thereby eliminating the interaction term from the regression. If that omitted term is positively correlated with $OPEN$, and if $\beta_2 < 0$, as predicted by theory, then OLS estimation of (6) will yield a fitted value of β_1 that is biased towards zero.

In order to examine the impact of openness on the slope of the Phillips curve after controlling for cross-country differences in the exchange rate regime, we construct an empirical counterpart to the variable EX. First, we take monthly data on the nominal effective exchange rate of country i over the same period as that used to measure PC_i. We then scale the exchange rate series by its mean, regress it on a constant and a time trend and calculate the residual standard error. These measures of exchange rate volatility are graphed in descending order in Appendix D. We identify three sub-groups within the sample, corresponding to high, intermediate and low levels of exchange rate volatility, and used this sample split as the basis for an exchange rate regime indicator, e, where $e_i = 2$ for the high levels of exchange rate volatility, $e_i = 1$ for intermediate levels of exchange rate volatility and $e_i = 0$ for low levels of exchange rate volatility (see Appendix D 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), Spain and France (who remained a part of the EMS only through widening the target zones for their currencies) and 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, which have not participated in a scheme like the EMS. The main exceptions to such rules are Greece (which appears in the floating group rather than the semi-fixed group) and Canada (which is in the semi-fixed group rather than the floating group).

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

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

where e^* denotes the mean of e_i.

4 Empirical results

In this section we investigate whether the lack of a correlation between openness and the slope of the Phillips curve reported in Temple (2002) is robust to using the PC index instead of the BMR index, and to controlling for an interaction between openness and the exchange rate regime. 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.[10] As the regressand is always a derived variable, t-ratios are calculated using the heteroscedasticity consistent standard errors due to White (1980). The absolute values of those t-ratios are reported in parentheses in Table 1. Openness to trade is measured as the mean of the ratio of total import spending to nominal GDP in country i over the time period used in measuring PC_i. The results of a chi-square test for residual normality due to Doornik and Hansen (1994) are also quoted (the null hypothesis is that the residuals are normally distributed).[11]

[10] We use the negatives of the BMR tradeoff measures in order to ensure that, like PC, the indices increase with the slope of the Phillips curve. Strictly speaking, one should use the value $\frac{(1-\pi_\bullet)}{\pi_\bullet}$ when making comparisons with PC. However, as the BMR parameter is actually negative for some countries, this transformation is non-monotonic.

[11] All regression estimates reported in this paper were obtained using the *PcGIVE* package of Hendry and Doornik (2001).

Table 1: Openness and the output-inflation tradeoff			
Regression	(1)	(2)	(3)
Dependent Variable^A	$-\pi_1$	$-\pi_2$	$-\pi_2$
Sample Size	19	19	19
$CONSTANT$	-.2825 (2.17)^B	-.5686 (4.46)	-.7187 (3.06)
$OPEN$	-.1395 (.39)	.3107 (.93)	.9114 (1.09)
$OPEN * EX$			-.4053 (.83)
Normality Test^C	1.54 (p = .46)	.30 (p = .86)	.35 (p = .84)
R^2	.004	.02	.06
Regression	(5)	(6)	(7)
Dependent Variable^A	$-\pi_2$	PC	PC
Sample Size	15	19	19
$CONSTANT$	-.7382 (3.79)	.5829 (3.32)	.0451 (.21)
$OPEN$	1.0584 (1.38)	-.4643 (1.11)	1.6888 (2.32)
$OPEN * EX$	-.5525 (1.28)		-1.4526 (3.60)
Normality Test^C	1.01 (p = .60)	5.58 (p = .06)	.85 (p = .65)
R^2	.18	.02	.33

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 new tradeoff parameter described in Section 3 of this paper.
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 finding in Temple (2002) that openness is both incorrectly signed and insignificant in a bivariate regression for the BMR measure of the output-inflation tradeoff. It is important to note that Temple's result is robust to restricting the sample from 42 countries to 19 countries.[12]

An obvious drawback to testing the Lane hypothesis using the 1948-86 BMR parameters is that the Bretton Woods fixed exchange rate system was effective for roughly two thirds of that period. The relationship between openness to trade and the slope of the Phillips curve would not be expected to operate under fixed exchange rate conditions, 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

[12] The BMR sample actually has 18 countries in common with the sample studied in section 3, New Zealand being the country that was included in the latter sample but not the former. However, we are able to expand the sample to 19 countries in Table 1 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.

the exchange rate regime. The slope coefficients are correctly signed in this model and the t-ratios are larger than in (2), but the magnitudes of the estimated 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 from the sample the United Kingdom and Norway (the two countries that are closest to having vertical Phillips curves according to $-\pi_2$) and Germany and Denmark (the two countries that are closest to having horizontal Phillips curves according to $-\pi_2$). These four countries are at the extreme ends of the range of Phillips curve parameters generated by the π_2 index, and are therefore likely to be the countries for which the index generates the largest amount of measurement bias. The results, presented in column (4), indicate that the effect of openness is significant at the 20% level, suggesting that measurement bias may be obscuring the link between openness and the slope of the Phillips curve.

In columns (5) and (6) we use PC as the dependent variable. The key result apparent from these two regressions is 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 .02 to .33. The importance of the interaction effect is due to the fact that the sample includes 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 adjustment relative to major trading partners. This means that large changes in import prices do not occur in those countries following expansions of the output gap, and that, as a result, Phillips curves in those countries have not been as steep as their openness to trade would predict.[13]

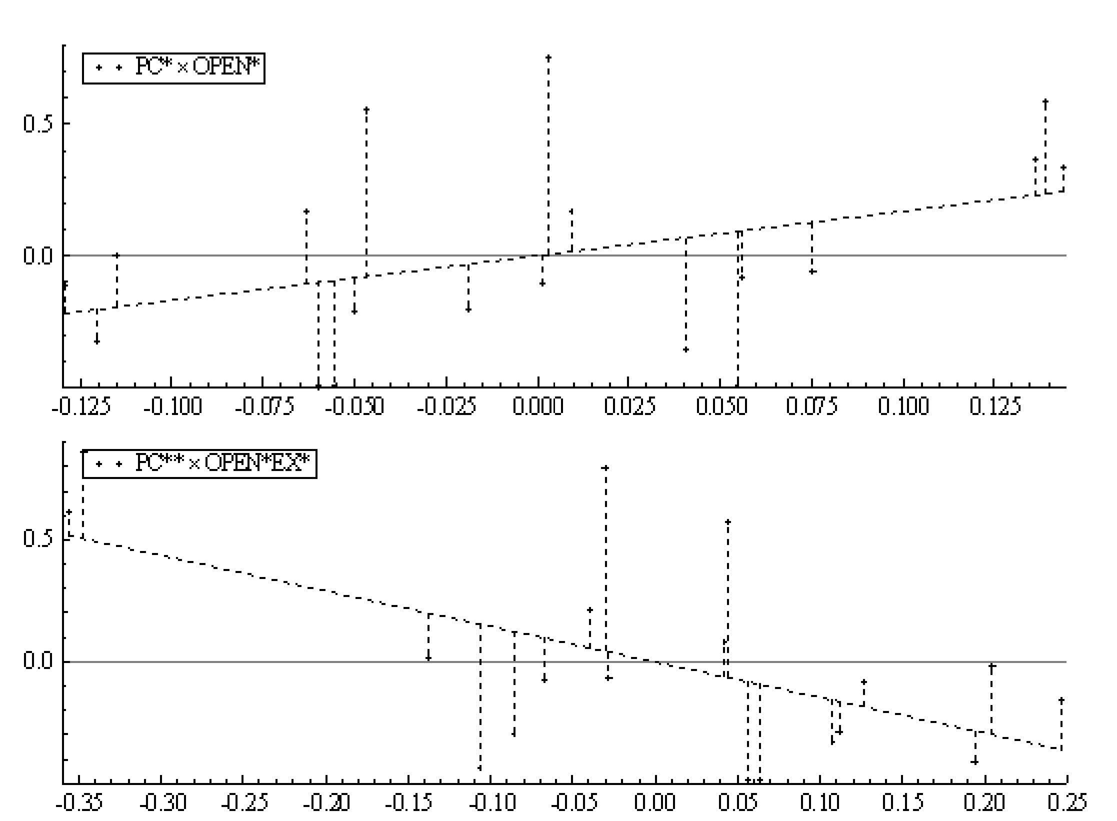
In section 3 we noted that the results of the exchange rate regime classification were slightly surprising, in that Canada was placed in the semi-fixed group rather than the flexible group, and Greece was placed in the flexible group rather than the semi-fixed group. If EX is reconstructed based upon Canada being in the flexible group and Greece being in the semi-fixed group, the results (which are not reported in the Table 1) are slightly stronger than those in column (6). The coefficient on $OPEN$ rises to 1.99 (robust t-ratio is 2.44), while that on $OPEN * EX$ becomes -1.64 (robust t-ratio is -3.52). It therefore appears that the relationship between openness and the slope of the Phillips curve is not dependent on the precise classification of exchange rate regimes outlined in section 3.

In order to cast further light on the relationship between the slope of the Phillips curve, openness to trade and the exchange rate regime, we consider the conditional scatter plots in Figure 1. We define PC^* as the set of residuals from a regression of PC on $OPEN * EX$, while $OPEN^*$ is the residual series from regressing $OPEN$ on $OPEN * EX$. The plot of PC^* against $OPEN^*$ and the associated line of best fit indicates the strength of the relationship between

[13] The coefficient estimates in column (6) indicate that fixing the exchange rate to the extent that countries such as Austria have done almost completely 'turns off' the effect of openness on the slope of the slope of the Phillips curve. However, given the large error bands associated with the coefficient estimates in Table 1, this conclusion can only be a tentative one.

the slope of the Phillips curve and trade openness after controlling for the exchange rate regime. Similarly, PC^{**} and $OPEN * EX^{*}$ are defined as the residual series from regressing PC and $OPEN * EX$ respectively on $OPEN$.

Figure 1: Conditional effects on the slope of the Phillips curve



The plots in Figure 1 suggest that the relationship between openness, the exchange rate regime and the slope of the Phillips curve may derive from the influence of a small number of observations in the sample. Specifically, the two observations in the northwest of the lower plot appear to explain a large part of the effect associated with the exchange rate regime in column (6). These points correspond to the observations for Greece and New Zealand, two of the countries from the flexible exchange rate group. Excluding these two countries from the sample causes the relationship in column (6) to disappear - the coefficient on openness falls to 0.4424 (robust t-ratio is 0.49), while the coefficient on the interaction term falls to -.6042 (robust t-ratio is 1.03). This finding reflects the fact that the relationship between openness and the slope of the Phillips curve is only apparent amongst those countries that have maintained flexible exchange rate regimes. The correlation between PC and $OPEN$ is .93 for the flexible exchange rate countries, but -.05 for the semi-fixed group and -.08 for the fixed group. As Greece and New Zealand represent the more open economies amongst the flexible exchange rate group (Australia, Japan and the United States are less open), deleting them from the sample means that a positive and significant relation between openness and the slope of the Phillips curve cannot be identified, essentially due to a lack of variation in the data.[14]

[14] If one adopts the alternative exchange rate regime classification in which Canada is in the flexible group and Greece is in the semi-fixed group, then the relationship between openness and the slope of the Phillips curve

In view of this finding, we suggest the following summary of the relationship between the slope of the Phillips curve, trade openness and the exchange rate regime in OECD countries. First, as approximately one quarter of OECD countries have maintained fixed exchange rate regimes since the 1970s, the conditions necessary for a correlation between openness and the slope of the Phillips curve have not been in place. Therefore, as expected, the slope of the Phillips curve does not increase with openness within that group of countries. Second, amongst those countries have followed semi-independent monetary policy, the relationship between trade openness and the slope of the Phillips curve is again absent. One reason for this may be that some measurement errors continue to affect estimates of the slope of the Phillips curve, e.g. due to fiscal policy affecting the output gap, or due to the controls used in (3)-(5) not handling the effects of all supply shocks. As the relationship between openness and the slope of the Phillips curve is likely to be a rather weak one amongst the semi-fixed exchange rate group, it may be obscured by small measurement errors. Third, amongst those countries that have maintained flexible exchange rate regimes the relationship between openness and the slope of the Phillips curve is positive and in line with the predictions based on the models of Romer (1993) and Lane (1997). In the cross-country regressions reported in this paper, it is the flexible exchange rate countries that drive the results, and as this group comprises just one quarter of the full sample, the regression results turn are sensitive to excluding a small number of countries from the sample. Thus, overall, empirical evidence on the Romer-Lane hypothesis based upon new measures of the slope of the Phillips curve is mixed - a positive relationship between openness and the slope of the Phillips curve is apparent amongst flexible exchange rate countries, but it is not apparent amongst countries that have sought to limit exchange rate fluctuations, suggesting that the underlying mechanism may be quite weak.

4.1 Controlling for closed economy effects on the Phillips curve

In this sub-section we extend the cross-sectional regressions in Table 1 to include further potential determinants of the slope of the Phillips curve. The first variable that we add to the analysis is a measure of inflation performance, calculated as the mean quarterly percentage inflation rate in a particular country over the period for the which the slope of its Phillips curve was estimated in section 3. The motivation for the inclusion of this regressor is that firms in high inflation countries have an incentive to reset prices more frequently than firms in low inflation countries, because high inflation means that firms face large relative price distortions if they leave absolute prices unchanged, see Ball, Mankiw and Romer (1988). We also condition on a set of labour market variables used by Temple (2002). $RIGIDITY$ measures the mean lag for the effect of unemployment on wages, and is taken from Grubb, Jackman and Layard (1983). The index decreases with the speed of wage adjustment and is therefore expected to enter the model with a negative sign. $INDEXATION$ takes the value $0, 1$ or 2 if wage indexation is, respectively, totally absent, partial or widespread. DUR measures the duration of price contracts and is also

re-appears, essentially because Canada is a relatively open economy within the flexible exchange rate group and has a steep Phillips curve relative to the United States and Japan. However, given that the data firmly suggest that Canada belongs in the semi-fixed exchange rate group (cf. the discussion in section 3), we do not attribute too much importance to this finding.

set to $0, 1$ or 2, with higher values indicating relatively short price contracts. Both variables are taken from Bruno and Sachs (1985) and are expected to enter the regression with a positive sign.[15] As observations on these variables are not available for all countries, the sample size changes slightly across model specifications. For clarity, the exact sample size is quoted above each set of results in Table 2.

Table 2: Closed economy effects on the slope of the Phillips curve					
Regression	(1)	(2)	(3)	(4)	(5)
Regressand	PC	PC	PC	PC	PC
Sample Size	19	19	18	17	17
$CONSTANT$	.36 (3.35)	.04 (.17)	.10 (.34)	.15 (.62)	.12 (.41)
$OPEN$		1.77 (1.81)	1.50 (1.42)	1.46 (1.12)	1.30 (1.46)
$OPEN * EX$		-1.52 (2.47)	-1.34 (2.10)	-1.35 (1.89)	-1.30 (2.17)
INF^2	.03 (2.87)	-.005 (.23)			
$RIGIDITY$			-.02 (.25)		
$INDEX$				-.02 (-.18)	
DUR					.03 (.31)
Normality	3.85(p=.15)	.74(p=.69)	1.08(p=.58)	.96(p=.62)	.94(p=.63)
R^2	.09	.33	.24	.26	.25
Notes: See notes B and C to Table 1.					

In column (1) of Table 2 we report a bivariate regression of PC on the square of mean inflation (the square of inflation is used instead of the level on grounds of best fit). The results are consistent with the notion that high inflation induces more frequent price-setting and a steepening of the Phillips curve. This is in line with the results obtained by BMR themselves using their single equation measure of the slope of the Phillips curve. The picture changes in regression (2), which adds open economy variables. The inflation term is insignificant and incorrectly signed, whilst openness and the interaction between openness and the exchange rate regime have the expected sign. We interpret this outcome as a result of the high degree of intercorrelation between openness, the exchange rate regime indicator and average inflation. As noted by Romer (1993), openness and the exchange rate regime tend to be important determinants of a country's inflation performance. Consequently, when average inflation is added to a regression for the Phillips curve parameter that already controls for open economy variables, it does not increase the explanatory power of the model.

The models in regressions (3)-(5) show that labour market variables do not help to explain cross-country differences in the slope of the Phillips curve. In each case the size of the coefficient on openness decreases slightly compared to that in column (6) of Table 1, and the standard errors increase quite sharply. However, this is mainly due to the reduced number of degrees of freedom compared to the Table 1 regressions.

[15] It should be noted that the labour market variables are measured over different periods to PC. As such, they are not suitable regressors in models for PC. We include them nevertheless, on the grounds that they are used in Temple (2002).

4.2 Controlling for potential regressor endogeneity

The relationship between openness, the exchange rate regime and the slope of the Phillips curve may be driven by reverse causation bias. For example, suppose a country faces a relatively flat short-run Phillips curve. A given sequence of aggregate demand shocks in this country will generate a relatively low variance inflation process and therefore a low variance detrended exchange rate, such that PC and $OPEN*EX$ could correlate negatively even when the Romer-Lane mechanism does not play a part in the determination of the output-inflation tradeoff.

In order to deal with potential endogeneity biases 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 $OPEN * EX$.[16] We maintain the assumption that $OPEN$ is exogenous, but now treat $OPEN*EX$ as endogenous, and draw instruments from the following set of variables: population size in 1990 (POP), the level and square of 1980 per capita income in US$ ($INCOME$) and an index of central bank independence (CBI) that decreases as a central bank becomes more independent.[17] In Table 3 we report regression estimates based upon different combinations of instruments. The absolute t-ratios given in parentheses are based on the corrected standard errors computed by the *PcGIVE* package, see Hendry and Doornik (2001). The R^2 statistic is not uniquely defined for 2SLS estimates, so here we report the regression standard error as a measure of fit for each specification. The Sargan statistic that we report tests the null hypothesis that the instrument set is valid in the sense that the instruments are uncorrelated with the errors generated from a regression of PC on all of the exogenous variables directly, see Hendry and Doornik (2001).

Table 3: Regressions for the tradeoff parameter estimated by 2SLS

Regression	$(1)^A$	$(2)^B$
Dependent Variable	PC	PC
Sample Size	19	19
$CONSTANT$	-.09 (.23)	.11 (.28)
$OPEN$	2.22 (1.56)	1.44 (.99)
$OPEN*EX$	-1.81 (2.13)	-1.28 (1.46)
Normality Test	.68 (p = .71)	.97 (p = .62)
Standard Error	0.38	0.38
Sargan Test	.19 (p = .66)	3.99 (p = .14)

The t-ratios are based on standard errors corrected for 2SLS estimation.
The null hypothesis for the Sargan test is that the instruments are uncorrelated with the errors in the unrestricted reduced form equation.
A. Instruments: $OPEN$, POP, CBI.
B. Instruments: $OPEN$, POP, $INCOME$, $INCOME^2$.

In column (1) we use population size, central bank independence and openness as instruments

[16] We also applied 2SLS to a model in which mean inflation is used as an explanatory variable. The results indicated that the lack of significance of the inflation term in Table 2 is robust to 2SLS estimation.

[17] The POP statistics are taken from the *International Financial Statistics* database maintained by the IMF, while $INCOME$ and CBI are taken from Romer (1993).

for the interaction between openness and the exchange rate regime. Population size is an effective instrument for the term $OPEN$ in $OPEN * EX$. It is more difficult to identify informative instruments for the exchange rate regime indicator, EX. One possibility is that conservative central banks are more likely to enforce fixed exchange rate regimes, and we therefore include CBI in the instrument. Of course, CBI may itself be endogenous, but we believe that it is much less likely to be so than the exchange rate regime indicator - the Sargan test outcome suggests that CBI is a valid instrument, though it should be noted that the small sample size used here may distort inferences based upon that procedure. The results in column (1) indicate that the effects of openness and the exchange rate regime on the slope of the Phillips curve increase relative to the OLS estimates in column (6) in Table 1, though due to the increased uncertainty in the estimation the coefficient on $OPEN$ actually loses significance. Still, we interpret this model as evidence that the relationship between the slope of the Phillips curve, trade openness and the exchange rate regime is not driven by endogeneity bias. In order to check the robustness of this finding, in estimating model (2) we delete CBI from the instrument set and add $INCOME$ and its square, the idea being that the choice of exchange rate regime may be income related. The coefficient estimates are of the same magnitude as those in column (6) of Table 1 (the OLS estimates), but due to an increase in estimation uncertainty the statistical significance of the effects decreases. We interpret this evidence as a symptom of the fact that it is difficult to identify reliable instruments for the exchange rate regime rather than a sign that past results were driven by endogeneity bias.

5 Summary

This paper has examined empirical evidence on the relationship between openness to trade and the slope of the Phillips curve. The importance of controlling for supply-side influences on output and inflation when measuring the slope of the Phillips curve, and of accounting for fixed exchange rate regimes in testing the implications of the theoretical models was emphasised. Results obtained for a sample of 19 countries indicated that greater openness to trade increases the slope of a country's short-run 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 are associated with the changes in import prices necessary to accelerate inflation adjustment. The result is consistent with the models of the output-inflation tradeoff in Romer (1993) and Lane (1997).

The robustness of the evidence was considered in some detail. It was shown that the correlation between openness and the slope of the Phillips curve derives from those countries that have followed flexible exchange rate regimes. Consequently the strength of the relationship estimated from a sample comprising fixed, semi-fixed and flexible exchange rate countries is sensitive to excluding certain countries from the analysis. We therefore interpret our results as preliminary support for the Romer-Lane hypothesis - more concrete evidence concerning this theoretical prediction can only be obtained through studying the Phillips curve in a larger sample of flexible exchange rate countries.

Appendix A: An open economy model of the Phillips curve (Romer (1993))

This appendix reviews the theoretical basis for the result in Romer (1993) that the slope of the Phillips curve is related to openness to trade. Romer considers a country that imports fraction a of the goods that its citizens consume. If e is the change from the preceding period in the log exchange rate, p^* the change in the log price index for foreign goods in foreign currency units, and p the change in the log price index for domestically produced goods in domestic currency units. Then the rate of consumer price inflation, x, is given by

$$x = a(e + p^*) + (1 - a)p$$

Romer then assumes that an individual's utility from consumption is a CES combination of his or her consumptions of different goods, with $\alpha < 1$ denoting the inverse of the elasticity of substitution between any two goods. Given that goods produced at home and abroad are imperfect substitutes in consumption, an expansion of domestic output drives down the relative price of domestically produced goods:

$$e + p^* - p = \alpha(y - y^*)$$

where y is the change in log domestic output and y^* the change in log foreign output.

Assuming that fraction f of domestic prices are flexible in the short-run and the remaining $1 - f$ are rigid, the inflation rate for domestically produced goods is

$$p = fp' + (1 - f)p^{\otimes}$$

where p' and $p^{\bullet}$ are the rates of inflation of prices that are flexible in the short-run and those that are fixed, respectively.

On the supply-side, it is assumed that flexible price inflation relative to consumer price inflation is an increasing function of output. If prices are initially at their equilibrium values then we have

$$p' - x = \phi y$$

Finally, money demand is given by

$$m - p = y$$

where m is the change in the log money stock. Analogous equations describe the rest of the world, which for simplicity consists of a single country. Letting an asterisk denote a foreign variable:

$$x^* = ap^* + (1 - a)(p - e)$$

$$p^* = fp'^* + (1 - f)p^{\otimes *}$$

$$p'^* - x^* = \phi y^*$$

$$m^* - p^* = y^*$$

Given these behavioural relations the effects of an increase in the money supply on output, domestic inflation and CPI inflation are given by

$$\frac{dy}{dm} = \frac{(1-f)\left[(1-f) + [(1-a)\alpha + \phi]f\right]}{\Delta}$$

$$\frac{dp}{dm} = \frac{f[(1-f)\phi + f(\phi+\alpha)\phi + (1-f)a\alpha]}{\Delta}$$

$$\frac{dx}{dm} = \frac{\phi f[(1-f) + f(\phi+\alpha)] + (1-f)a\alpha(1+f\phi)}{\Delta}$$

where

$$\Delta \equiv [(1-f) + \phi f][(1-f) + (\phi+\alpha)f]$$

It can be seen that the effect of a monetary expansion on output is smaller in a more open economy, and that its effects on both domestic and CPI inflation are larger. Thus, the output-inflation tradeoff, $\frac{dy}{dx}$ is less favourable in a more open economy.

Appendix B: Variable definitions and sample periods

The variables used in equations (3)-(5) are defined below. All variables are in seasonally adjusted form, and, unless otherwise stated, refer to natural logarithms of the variables measured. The data sources are the *OECD quarterly national accounts* and the IMF's *International Financial Statistics*, unless otherwise stated.

p is the consumer price index (CPI). The CPI excludes mortgage interest payments in all countries except Australia, New Zealand, the United Kingdom and the United States (pre-1983 in the US case). In order to ensure the comparability of price data across countries, we use the implicit consumption deflator as the measure of prices for those four countries.

ulc is average unit labour costs for the whole economy and is constructed as follows:

$$\textit{unit labour costs} = \textit{total wages and salaries - constant price GDP}$$

usoil is the US$ price of a barrel of crude oil. These data were supplied by John Muellbauer.

gap measures the deviation of constant price GDP, y_t, from a stochastic trend. This measure of the output gap is closely related to that in Aron and Muellbauer (2000) and draws on the *STAMP* computer package of Koopman, Harvey, Doornik and Shephard (1995). Log income, 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

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

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

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

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

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 v_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 $c + v\chi_t$, and the output gap is measured as $y_t - c - v\chi_t$. 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.

Sample periods

The sample periods used to fit equations (3)-(5) were as follows: Australia: 1976q1-1997q3. Austria: 1976q1-1995q4. Belgium: 1981q2-1997q3. Canada: 1978q2-1997q1. Denmark: 1978q1-1993q4. Finland: 1976q2-1997q3. France: 1979q2-1994q2. Germany: 1976q1-1997q3. Greece: 1981q2-1991q1. Italy: 1976q1-1996q3. Japan: 1979q1-1997q3. Netherlands: 1978q2-1997q3. New Zealand: 1981q4-1997q3. Norway: 1979q1-1997q3. Spain: 1981q2-1996q4. Sweden: 1976q1-1997q2. Switzerland: 1976q2-1997q3. United Kingdom: 1976q1-1997q3. United States: 1976q1-1997q3.

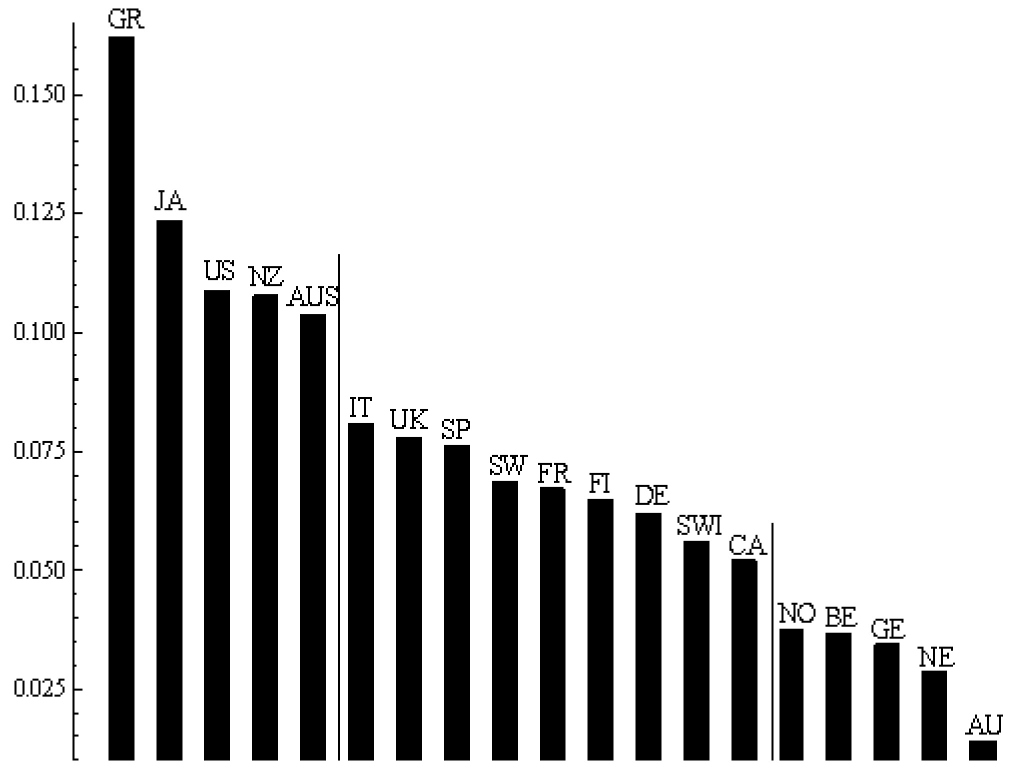
Appendix C: The PC index

The measures of the slope of the Phillips curve, PC, that we obtain using the methods described in section 3 are as follows: Australia (.4287), Austria (0), Belgium (.3426), Canada (1.2497), Denmark (0), Finland (.6671), France (1.0485), Germany (.3037), Greece (1.0600), Italy (0), Japan (.4538), Netherlands (0), New Zealand (1.3095), Norway (.4077), Spain (0), Sweden (.1381), Switzerland (.4375), United Kingdom (.3922), United States (.2477).

Appendix D: Notes on the construction of the EX dummy

Figure 2 presents measures of the volatility of linearly detrended nominal effective exchange rate data for 19 countries. The abbreviations used 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. 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 2. 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 and the United States (the flexible exchange rate group).

Figure 2: Detrended exchange rate volatilities



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