Marginal Utility and International Relative Prices

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Abstract

A wide range of international macroeconomic models link the real exchange rate to a ratio of marginal utilities. These include models based on non-traded goods or those with a variety of frictions including trading costs and pricing to market. Empirical work has found little evidence of this link. We augment the measurement of international relative prices to allow for incomplete pass-through of nominal exchange rate changes to domestic inflation. With that realistic feature, there is a clear connection between relative prices and relative marginal utilities in a panel of OECD countries since 1960. Preference parameters are significant and consistent with theory.

Keywords: real exchange rate, consumption, marginal utility

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1. Introduction

A large volume of research in international macroeconomics focuses on explanations for the persistence in real exchange rates. But most models of persistent deviations from purchasing power parity imply that relative marginal utilities of consumption across countries should be as persistent as the real exchange rate. Unfortunately, there is little evidence of a link between the real exchange rate and relative consumption. Obstfeld and Rogoff (2000b) list this among the key, unresolved issues in open economy macroeconomics.

This paper first documents the generality of the marginal-utility condition to various international dynamic equilibrium models. Second, the condition is studied in historical data. Once we allow for, and measure, incomplete pass-through of nominal exchange rate changes to domestic prices, we find a significant link between relative prices and relative marginal utilities.

Section 2 provides background on the condition we test. Section 3 outlines a parametric model of utility, allowing a role for government spending. Section 4 gives statistical evidence for a panel of OECD countries since 1960. Section 5 extends the utility model to allow a role for leisure or real balances, and tests this. Section 6 concludes.

2. Background

Backus and Smith (1993) and Kollmann (1995) observed that a range of international macroeconomic models with nontraded goods link the real exchange rate to relative consumption. For example, suppose that two countries i and j have consumptions c_i and c_j and price levels p_i and p_j. If the nominal exchange rate between them is e_{ij} and the two countries have power utility with exponent α then:

$$\frac{p_{it}}{p_{jt}e_{ijt}} = \left(\frac{c_{jt}}{c_{it}}\right)^{\alpha}. \tag{1}$$

The strong empirical persistence of the real exchange rate requires a corresponding autocorrelation in the relative consumption ratio. However, Backus and Smith (1993) found that in OECD data, the consumption ratio was negatively autocorrelated in 27 country pairs, whereas the real exchange rate was positively autocorrelated in 28 country pairs.

Further, they found that the cross correlation of the consumption ratio and the real exchange rate over their entire sample was only .045. Kollmann (1995) studied the link (1) in levels as well as growth rates, with equally negative findings.

Evidence on international pricing suggests that departures from the law of one price, rather than traditional non-traded goods, account for most movements in real exchange rates. Engel (1993), Goldberg and Knetter (1997) and Engel and Rogers (1997) document this fact. In response to this evidence, a growing segment of the literature has focused on the international pricing behaviour of imperfectly-competitive firms. "Pricing to market" (PTM), first suggested by Krugman (1987) and Dornbusch (1987), supposes that the prohibitive cost of goods arbitrage allows firms to price discriminate across internationally segmented markets. If prices are sticky, then nominal exchange rate movements will not be passed through into goods prices. The resulting deviation from the law of one price causes a corresponding movement in the real exchange rate. Betts and Devereux (1996), Kollman (1996) and Chari, Kehoe and McGratten (1997, 1998), and Bergin and Feenstra (1999) have built dynamic general equilibrium models to investigate the quantitative impact of PTM. Although these models have enjoyed some success, they often predict counterfactual increases in international consumption persistence.

In response to the expanding PTM literature, Obstfeld and Rogoff (2000a) argue that the PTM paradigm is inconsistent with the empirical evidence. First, long periods of sticky prices are not as empirically plausible as sticky wages. Second, they argue that the evidence of deviations from the law of one price is based on aggregated price data. As a result, the role of nontradable costs may exert a stronger impact for real-exchange rate movements than currently believed. Further, the price of these nontradable costs such as rent, distribution and advertising may deviate sharply across countries. Finally, they cite the inconsistency of pricing-to-market models with international expenditure switching following nominal exchange rate movements. Expenditure switching does not occur in PTM models because the price of exports and imports are expressed in local currency so that households are insulated from relative price effects following exchange-rate adjustments. Following these concerns regarding the PTM paradigm, Obstfeld and Rogoff (2000a) developed a model with nontraded goods and sticky wages. However, this model still links the real exchange rate to relative marginal utilities.

The same is true of models with explicit trading costs. Dumas (1992) was able to augment the persistence of the real exchange rate in an environment characterized with spatially separated goods markets. In his model, a shipping cost is required in order to transfer goods across markets. In this manner, households are limited to the set of goods that are physically located within their own market. Consequently, a cone of "no shipping" arises. Prices and marginal utilities can diverge within this cone. Obstfeld and Rogoff (2000b) suggested that trade costs can reconcile models with a wide range of empirical puzzles in international macroeconomics. But Engel (2000) observed that this approach too links prices and marginal utilities (consumption) in an unrealistic way.

Thus a wide range of models – including nontraded goods, pricing to market, sticky prices or wages, or transport costs – runs afoul of the condition linking relative prices and relative marginal utilities. In this paper we reconsider the parametric models of utility used in past studies of this condition. We take advantage of its static nature to estimate preference parameters without a complete model. We also explore whether the measurement of relative prices affects the empirical findings.

3. Utility Model

Utility in country i is of the power form:

$$u(x_{it}) = \begin{cases} x_{it}^{1-\alpha}/(1-\alpha) & \alpha > 0, \quad \alpha \neq 1; \\ \ln x_{it} & \alpha = 1. \end{cases} \tag{2}$$

where x_{it} is an aggregator over private consumption c_{it} and government consumption g_{it}. This aggregate in turn is of the CES form:

$$x_{it} = [\mu c_{it}^{\omega} + (1-\mu) g_{it}^{\omega}]^{1/\omega} \tag{3}$$

We test two special cases of this aggregator. In the first, $\omega = 0$, which yields the Cobb-Douglas case:

$$x_{it} = c_{it}^{\mu} g_{it}^{1-\mu}. \tag{4}$$

In the second special case, $\mu = 1$ so that public expenditure does not directly affect utility and $x_{it} = c_{it}$.

The condition we test is that international relative prices are related to the ratio of marginal utilities of consumption:

$$\frac{p_{it}}{p_{jt} e_{ijt}^{\gamma}} = \frac{u_c(x_{it})}{u_c(x_{jt})}. \tag{5}$$

The right-hand side is the real exchange rate if $\gamma = 1$. But a novel feature here is that we do not assume that the law of one price holds. To allow for incomplete pass-through of nominal exchange rate changes to domestic prices, we estimate γ along with the parameters of marginal utility. In contrast with microeconomic measures of pass-through, we estimate the γ that leads to the best fit of the link (5) between prices and relative marginal utilities. The underlying question is: Does *any* aggregate measure of international relative prices (even if it is not the real exchange rate) have statistical properties like those of international relative marginal utility?

4. Statistical Results

Data are for a panel of nine countries: Canada, Denmark, Finland, France, Italy, Japan, Sweden, the UK, and the US. The panel was selected based on the availability of measures of private consumption excluding durables. For Canada, the UK, and the US, the data run from the 1960s, for France, Italy, and Japan from the 1970s, and for the Scandinavian countries from 1988. The appendix provides exact data definitions and sources.

This version of the paper uses a balanced panel from 1970:I to 1998:III, which omits the three Scandinavian countries. This gives 112 observations for each of the six remaining countries. The next version will include statistics for unbalanced panels and hence use more observation-country pairs.

Given the parametric utility model (2)–(3), we estimate the equilibrium condition (5) in growth rates, to allow the use of standard tools of inference. That implies:

$$\begin{aligned}\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma \ln e_{ijt} = \kappa_{ij} &+ \frac{1-\alpha-\omega}{\omega} \Delta \ln\bigl[\mu (c_{it})^{\omega} + (1-\mu)(g_{it})^{\omega}\bigr] \\ &- \frac{1-\alpha-\omega}{\omega} \Delta \ln\bigl[\mu (c_{jt})^{\omega} + (1-\mu)(g_{jt})^{\omega}\bigr] \\ &+ (\omega - 1)\bigl[\Delta \ln(c_{it}) - \Delta \ln(c_{jt})\bigr]\end{aligned} \tag{6}$$

where κ_{ij} is an intercept term. Equation (6) describes one of five regression equations that are estimated as a nonlinear system. The base country in each regression equation is the US, so there is potential for cross-sectional dependence of the type highlighted by O'Connell (1998). For that reason, estimation is by generalized least squares. Cross-equation restrictions are imposed to increase the power of the regression, and allow sufficient moment conditions to obtain over-identification. As the notation shows, the restrictions are that the preference parameters — α, μ, and ω — are common across countries. We also estimate γ, the degree of pass through of exchange rate changes. We later test the stability of these parameters across countries.

The Cobb-Douglas special case gives:

$$\begin{aligned} \Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma \ln e_{ijt} = \kappa_{ij} &+ (\mu - 1 - \alpha\mu)\big[\Delta \ln c_{it} - \Delta \ln c_{jt}\big] \\ &+ (1 - \mu - \alpha + \alpha\mu)\big[\Delta \ln g_{it} - \Delta \ln g_{jt}\big]. \end{aligned} \tag{7}$$

When utility depends only on private consumption, $\mu = 1$ so:

$$\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma \ln e_{ijt} = \kappa_{ij} - \alpha\big[\Delta \ln c_{it} - \Delta \ln c_{jt}\big]. \tag{8}$$

Our findings begin with Table 1, which shows the growth-rate version of the relationship studied by Backus and Smith (1993) and Kollmann (1995). The dependent variable is the growth rate of the real exchange rate. There is no statistical relationship between this measure and the growth of relative marginal utility on the right-hand side of the system of equations. The explanatory power is zero, and $\hat{\alpha}$, though positive, is statistically insignificant.

Table 2 instead measures international relative prices using a weight γ on the nominal exchange rate which is estimated and so can differ from 1. The second column gives the CES estimates, and shows that ω cannot be distinguished from zero. The asymptotic p-value for this test is 0.26. Then the third column shows that the Cobb-Douglas case including government spending also can be specialized to include only private consumption. Testing that $\mu = 1$ yields an asymptotic p-value of 0.98. The last column then focuses on the case with private consumption. Here $\hat{\alpha}$ is positive, as required by the theory, and significantly different from zero (with a p-value of 0.00).

The bad news for the theory is that the explanatory power or fit of the two sides of the equation is quite low. The last two rows of Table 1 show the range for $\bar{R}^2$ and the Durbin-Watson statistic. There generally is positive, residual autocorrelation (especially for Italy, though countries are not shown individually) and the relative inflation measure is much more volatile than the weighted relative consumption growth measure. Also, the intercepts, $\hat{\kappa}_{ij}$s, are not all zero; half of them differ from zero at conventional significance levels. This finding means that there is a drift in differential inflation that does not match the drift in differential consumption growth.

To summmarize, there is some evidence in support of the correlation. We find a positive, significant exponent in power utility, and credible estimates of other parameters of utility. Evidence against the equilibrium condition takes the form of intercepts, residual autocorrelation, and low explanatory power.

The measure of the pass-through of nominal exchange rate changes, $\hat{\gamma}$, is positive, with a p-value of 0.08. We also examine the effect of allowing this measure to vary across countries. In the case with private consumption only, the largest values for $\hat{\gamma}_{ij}$ (with standard error) are 0.03 (0.009) for Finland and 0.081 (0.024) for Sweden. Other findings are essentially unchanged.

To examine these findings in more detail, we next free up the parameters across countries. For each country we estimate α_{ij}, γ_{ij}, and κ_{ij} relative to country j, the U.S. Table 3 contains the results for the five countries included in the panel. For all 5, the hypothesis that $\mu = 1$ could not be rejected, and so the table shows only the results with private consumption. Tables 2 and 3 can be combined to reject the hypothesis that the parameters are stable across countries. However, the findings are similar to those of Table 2: positive and significant utility parameters $\hat{\alpha}_i$s (except for Italy, where this is insignificant); small, positive, pass-through parameters $\hat{\gamma}_{ij}$; relatively low explanatory power; positive autocorrelation of the residuals.

Table 4 contains country-by-country estimates that use all the data and now include the Scandinavian countries. (This will be compared to the unbalanced panel, once it is estimated in the next version of this paper.) Countries are listed in decreasing order of sample size. Now there is somewhat less evidence of positive and significant $\hat{\alpha}_i$s, but the general picture remains the same.

5. Nonseparability of Leisure or Real Balances

Given the mixed results so far, we next study other models of utility that have been used in international macroeconomics. These models involve a nonseparability between consumption goods and either leisure or real money balances.

First, consider a nonseparability between consumption and leisure. The period utility function is again of power form:

$$u(z_{it}) = \begin{cases} z_{it}^{1-\alpha}/(1-\alpha) & \alpha > 0, \quad \alpha \neq 1; \\ \ln z_{it} & \alpha = 1. \end{cases} \tag{9}$$

where z_{it} combines the CES aggregate x_{it} (3) of private and public consumption with a measure of employment l_{it}:

$$z_{it} \equiv x_{it} - \delta l_{it}^{\eta}. \tag{10}$$

This form is chosen to nest the case studied by Greenwood, Hercowitz and Huffman (1988), Hercowitz and Sampson (1991), Devereux, Gregory, and Smith (1991) and Correia, Neves, and Rebelo (1995). Those authors set $\alpha = 1$ and $x_{it} = c_{it}$ for analytical tractability. Because we are not solving a model, we consider a more general form.

As an example, consider the case in which $x_{it} = c_{it}$, as we found in section 4. Then the estimating equation is:

$$\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma \Delta \ln e_{ijt} = \kappa_{ij} - \alpha \big[\Delta \ln(c_{it} - \delta l_{it}^{\eta}) - \Delta \ln(c_{jt} - \delta l_{jt}^{\eta})\big]. \tag{11}$$

The relative growth of marginal utility across countries now contains labour supply measures. International employment differences tend to be positively autocorrelated, so this addition to the statistical model may produce persistence that more closely matches the persistence in the relative inflation measure. (Recall that residual autocorrelation was high in several countries in section 4).

We estimated the panel (11), with l_{it} measured as the OECD's index of total employment. The results were entirely negative. The weight on employment, $\hat{\delta}$ had the wrong sign and $\hat{\alpha}$ was negative. The fit was not improved over that in section 4. We conclude that the time series correlation between relative employment and relative prices does not match the restrictions of this model of marginal utility.

A number of recent research papers in international finance have included real money balances in the utility function. We next follow Chari, Kehoe, and McGrattan (1997, 1998, 2000) who incorporate a nonseparability between real balances, m_{it}, and private consumption. Their utility function is:

$$U(c_{it}, m_{it}) \;=\; \frac{1}{1-\alpha}\left[\left(\mu\; c_{it}^{\omega} + (1-\mu)\; m_{it}^{\omega}\right)^{\frac{1}{\omega}}\right]^{1-\alpha}, \tag{12}$$

which gives rise to the following marginal utility with respect to consumption:

$$U_c \;=\; \mu c_{it}^{\omega-1}\left[\mu c_{it}^{\omega} + (1-\mu) m_{it}^{\omega}\right]^{\frac{1-\alpha-\omega}{\omega}}. \tag{8}$$

Real balances were measured as broad money (generally M2) divided by the consumption deflator. We estimated the relative inflation equations using both CES and Cobb-Douglas aggregators over private consumption and real balances. There is a small role for real balances, while the general picture is unchanged. The curvature of utility, $\hat{\alpha}$ is small and significant, as is the pass-through coefficient, $\hat{\gamma}$. Explanatory power, measured by $\bar{R}^2$ ranges from 0.06 to 0.31, and residuals are generally persistent. Including real money balances in our model of marginal utility does not lead to a large statistical improvement.

6. Conclusion

Several statistical refinements can be made to these preliminary results. Chief among these will be the use of unbalanced panels, so that all observations can be used.

Our economic findings raise two key questions. First, there is no evidence that relative marginal utility matches the properties of the real exchange rate. We have shown that this remains a stumbling block for many models that otherwise work well. There is some evidence that relative inflation (with a small role for the nominal exchange rate) does fit with the standard models of marginal utility. Our future work will focus on theoretical models which break the link between the real exchange rate and relative marginal utility.

Second, even with this change the fit of the static relationship is poor. We also hope to see whether we can generalize about the shocks under which the two sides of the equation move together and the shocks under which they do not.

Data Appendix

The data for government expenditure, g, private consumption, c, and the implicit price deflator of private consumption, p are from the OECD *Quarterly National Accounts*. Nominal exchange rates, e, come from the IMF's *International Financial Statistics* (IFS) database. Exchange rates are local currency prices of a US dollar. Private consumption is measured as the sum of consumer nondurables and services expenditure. Both private and government consumption are in constant 1992 dollars. The nominal exchange rate is the average rate in the quarter.

Labour data are non-seasonally adjusted index series with 1995 acting as the base year. These data are generally defined as "total employment" in each country, so that individuals who have worked for any duration over the reference period are included as employed. The quarterly series are collected from the *Main Economic Indicators*, published by the OECD.

Broad money data are also extracted from the *Main Economic Indicators*. These series are generally defined as M2, but for some countries the only available data are M3. Specifically, M2 is available for Canada, Denmark, Finland, France, Italy and the UK. Data for Japan are reported as M2 plus CDs. Finally, broad money is reported as M3 for the US, France and Sweden. Real money balances are defined as the nominal money supply in local currency divided by the consumption price deflator. All money data are seasonally unadjusted.

The panel includes Canada, Denmark, Finland, France, Italy, Japan, Sweden, the United Kingdom, and the United States. Graphs of the series were examined to identify possible problems with the data. Seasonal patterns in Swedish government spending and consumption, in Japanese consumption, and in the employment index and broad money were removed with the `ESMOOTH` command in Rats^{TM}. Table A1 gives the time span available for each time series.

Table A1: Availability of Time Series

	Canada	Denmark	Finland	France	Italy
c	61:I–99:I	88:I–99:I	88:I–99:I	70:I–98:3	70:I–98:3
g	61:I–99:I	88:I–99:I	88:I–99:I	70:I–98:3	70:I–98:3
p	61:I–99:I	88:I–99:I	88:I–99:I	70:I–98:3	70:I–98:3
e	55:I–99:I	55:I–99:I	71:I–99:I	55:I–99:I	55:I–99:I
l	60:I-99:1	80:I-98:IV	60:I-99:I	77:IV-99:I	60:I-99:I
M	68:I-98:IV	70:I-99:I	60:I-98:IV	77:IV-98:IV	74:IV-99:I

	Japan	Sweden	UK	US
c	70:I–98:I	80:I–98:4	63:I–99:I	59:I–99:I
g	55:2–99:I	80:I–98:4	55:I–99:I	59:I–99:I
p	55:2–99:I	80:I–98:4	55:2–99:I	59:I–99:I
e	57:I–99:I	55:I–99:I	55:I–99:I	–
l	60:I-99:I	61:2-99:I	60:I-99:I	60:I-99:II
M	60:I-99:I	60:I-99:I	82:II-99:I	60:I-99:I

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Table 1. Panel Estimates: Real Exchange Rate

$$\begin{aligned}\Delta \ln p_{it} - \Delta \ln p_{jt} - \Delta \ln e_{ijt} = \; & \kappa_{ij} + \frac{1-\alpha-\omega}{\omega} \; \Delta \ln\big[\mu(c_{it})^{\omega} + (1-\mu)(g_{it})^{\omega}\big] \\ & - \frac{1-\alpha-\omega}{\omega} \; \Delta \ln\big[\mu(c_{jt})^{\omega} + (1-\mu)(g_{jt})^{\omega}\big] \\ & + (\omega - 1) \; \big[\Delta \ln(c_{it}) - \Delta \ln(c_{jt})\big] + \epsilon_{ijt}\end{aligned}$$

Coefficient	CES case	Cobb-Douglas $\omega = 0$	Private c only $\omega = 0$ $\mu = 1$
$\hat{\omega}$	0.872 (0.255)		
$\hat{\mu}$	239571525.077 (17580040.079)	1.011 (0.102)	
$\hat{\alpha}$	0.115 (0.195)	0.113 (0.197)	0.105 (0.161)
$\bar{R}^2$ min/max	0.00/0.01	0.00/0.01	0.00/0.01
$\hat{\rho}_{\epsilon}$ max/min	0.33/0.18	0.33/0.18	0.33/0.18

Notes: Autocorrelation-consistent standard errors are in brackets. $\hat{\rho}_{\epsilon}$ is the residual, first-order autocorrelation coefficient. The United States is country j.

Table 2. Panel Estimates with Incomplete Pass Through

$$\begin{aligned}\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma_{ij} \Delta \ln e_{ijt} = \kappa_{ij} &+ \frac{1-\alpha-\omega}{\omega} \; \Delta \ln\big[\mu(c_{it})^{\omega} + (1-\mu)(g_{it})^{\omega}\big] \\ &- \frac{1-\alpha-\omega}{\omega} \; \Delta \ln\big[\mu(c_{jt})^{\omega} + (1-\mu)(g_{jt})^{\omega}\big] \\ &+ (\omega - 1) \; \big[\Delta \ln(c_{it}) - \Delta \ln(c_{jt})\big] + \epsilon_{ijt}\end{aligned}$$

	CES case	Cobb-Douglas $\omega = 0$	Private c only $\omega = 0$ $\mu = 1$
Coefficient			
$\hat{\omega}$	-3.261 (2.920)		
$\hat{\mu}$	1.000 (0.001)	1.001 (0.041)	
$\hat{\alpha}$	0.393 (0.040)	0.374 (0.040)	0.374 (0.039)
$\hat{\gamma}$	0.019 (0.012)	0.021 (0.012)	0.021 (0.012)
$\bar{R}^2$ min/max	-0.06/0.29	-0.06/0.29	-0.06/0.29
$\hat{\rho}_{\epsilon}$ max/min	0.84/0.47	0.84/0.47	0.84/0.47

Notes: Autocorrelation-consistent standard errors are in brackets. $\hat{\rho}_{\epsilon}$ is the residual, first-order autocorrelation coefficient. The United States is country j.

Table 3. Country-by-Country Estimates: 1970:I to 1998:III

$$\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma_{ij} \Delta \ln e_{ijt} = \kappa_{ij} - \alpha_{ij} \big(\Delta \ln c_{it} - \Delta \ln c_{jt} \big) + \epsilon_{ijt}$$

Country	$\hat{\alpha}_{ij}$	$\hat{\gamma}_{ij}$	$\hat{\kappa}_{ij} \cdot 100$	$\bar{R}^2$	$\hat{\omega}_{\epsilon}$
Canada	0.491	0.034	0.162	0.25	0.45
	(0.045)	(0.036)	(0.084)		
U.K.	0.629	-0.015	0.618	0.34	0.53
	(0.118)	(0.015)	(0.132)		
France	0.185	0.032	0.293	0.11	0.51
	(0.071)	(0.013)	(0.118)		
Italy	-0.077	0.054	1.105	0.14	0.79
	(0.094)	(0.018)	(0.150)		
Japan	0.287	0.001	-0.137	-0.01	0.53
	(0.149)	(0.017)	(0.190)		

Notes: Autocorrelation-consistent standard errors are in brackets. $\hat{\rho}_{\epsilon}$ is the residual, first-order autocorrelation coefficient. The United States is country j.

Table 4. Country-by-Country Estimates: All data and countries

$$\Delta \ln p_{it} - \Delta \ln p_{jt} - \gamma_{ij} \Delta \ln e_{ijt} = \kappa_{ij} - \alpha_{ij} \big(\Delta \ln c_{it} - \Delta \ln c_{jt}\big) + \epsilon_{ijt}$$

Country	Period	$\hat{\alpha}_{ij}$	$\hat{\gamma}_{ij}$	$\hat{\kappa}_{ij} \cdot 100$	$\bar{R}^2$	$\hat{\rho}_{\epsilon}$
Canada	1960:I – 1999:I	0.482	0.026	0.164	0.26	0.36
		(0.100)	(0.042)	(0.069)		
U.K	1963:I – 1999:I	0.510	0.005	0.485	0.25	0.51
		(0.113)	(0.015)	(0.109)		
France	1970:I – 1998:III	0.158	0.040	0.285	0.11	0.49
		(0.092)	(0.016)	(0.115)		
Italy	1970:I – 1998:I	-0.304	0.072	1.103	0.17	0.73
		(0.125)	(0.025)	(0.136)		
Japan	1970:I – 1998:I	0.068	0.005	-0.162	0.01	0.49
		(0.151)	(0.017)	(0.188)		
Sweden	1980:I – 1998:IV	-0.548	0.079	0.713	0.13	-0.02
		(0.429)	(0.020)	(0.297)		
Denmark	1988:I – 1999:I	0.027	0.010	-0.159	0.04	0.19
		(0.062)	(0.012)	(0.103)		
Finland	1988:I – 1999:I	0.007	0.032	-0.044	0.15	0.16
		(0.078)	(0.012)	(0.088)		

Notes: Autocorrelation-consistent standard errors are in brackets. $\hat{\rho}_{\epsilon}$ is the residual, first-order autocorrelation coefficient. The United States is country j.