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PRODUCTIVITY AND THE EURO-DOLLAR EXCHANGE RATE PUZZLE

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ABSTRACT

This paper documents the evidence for a productivity based model of the dollar/euro real exchange rate over the 1985-2001 period. We estimate cointegrating relationships between the real exchange rate, productivity, and the real price of oil using the Johansen (1988) and Stock-Watson (1993) procedures. We find that each percentage point in the US-Euro area productivity differential results in a five percentage point real appreciation of the dollar. This finding is robust to the estimation methodology, the variables included in the regression, and the sample period. We conjecture that productivity-based models cannot explain the observed patterns with the standard set of assumptions, and describe a case in which the model can be reconciled with the observed data.

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1 INTRODUCTION

A commonplace observation is that exchange rates appear to behave in seemingly inexplicable ways. Even by this standard, the euro's behavior has confounded most observers' expectations. Rather than fulfilling expectations of appreciation, the euro spent the better part of its first two years of existence depreciating against the dollar.

Many arguments have been forwarded for this trend. They include income surprises (as signals of revisions in trend productivity growth, in Corsetti and Pesenti, 2000); black market holdings of euros (Sinn and Westermann, 2001); overly restrictive fiscal policy (Cohen and Loisel, 2000); and higher equity returns on New Economy assets in the U.S. relative to the Euro Area, a view common in the financial sector.¹

While it is easy enough to point to trends in one's favorite variable, it is much more difficult to pinpoint a statistical correlation that stands up to close scrutiny. In this paper, we document a robust but inherently puzzling correlation, displayed below in Figure 1.

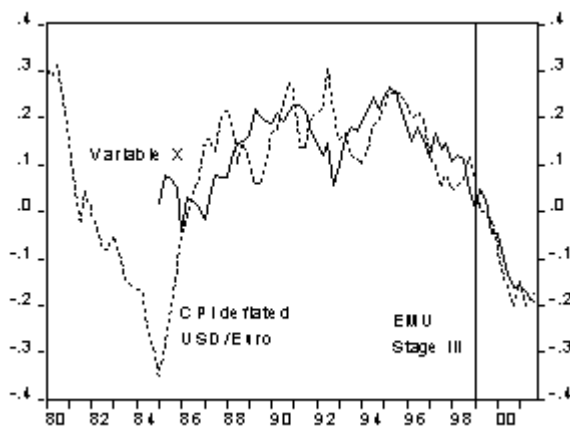


Figure 1: Real Dollar/Euro Rate and Variable X

¹ See for instance Rosenberg, 2000; Owen, 2001; van de Craen and Vanden Houte, 2001. The focus on capital flow composition is consistent with this perspective. See for instance Brooks et al. (2001).

With the exception of the mid-1980's surge in the dollar's value, variable X has exhibited a strong correlation with the synthetic real euro.

The mysterious "Variable X" in the figure is the Euro area-US log productivity differential, multiplied by 5.0. Below we document that this correlation is not "spurious" in a specific statistical sense; in other words the real exchange rate appears to be cointegrated with a broad productivity differential. Over a number of specifications, sample periods, and estimation techniques, one finds that each one percentage point increase in the U.S.-Euro area productivity differential results in between a four to five percent real appreciation in the dollar/euro exchange rate.

While this correlation accords with popular explanations linking the New Economy to asset markets, the magnitude of this correlation is much larger than what is to be expected from the dominant theoretical model for productivity effects, the Balassa-Samuelson model. Similarly, the neoclassical models of Lucas (1982) and Stockman (1980) that comprise the other major competing paradigm typically do not predict a real exchange rate appreciation in response to a productivity shock.

In this paper, we propose the following reconciliation of our findings with theory. Faster productivity growth can increase U.S. aggregate demand through two channels, consumption and investment. Forward looking consumers increase consumption due to expectations of higher future income and higher productivity increases expected future profits, raising investment demand (Bailey, Millard, and Wells, 2001). We focus on the former effect in Ricardian framework and argue that U.S. demand for domestic goods is so high due to wealth effects that it overwhelms the downward pressure higher productivity exerts upon the terms of trade.

Consequently, on net the real exchange rate appreciates.

This paper proceeds in the following manner. In Section 2, data preliminaries are discussed, and a possible theoretical explanation forwarded. In section 3, the econometric methodology is discussed. The empirical estimates and out-of-sample performance of the models are examined in Section 4. Some concluding remarks are included in Section 5.

2. Patterns and Hypotheses

The path of the real dollar/euro exchange rate is presented in Figure 2. Two measures are used as price deflators: the CPI and the PPI. Interpretation of these movements is complicated by the obvious fact that the euro is a recent innovation so that pre-1999 trends have to be proxied by movements in the synthetic euro. Abstracting from this complication, we observe that the dollar/euro exchange rate has been characterized by two large swings: the mid-1980's dollar spike and the current dollar surge. Over this long span, movements in the CPI or PPI deflated

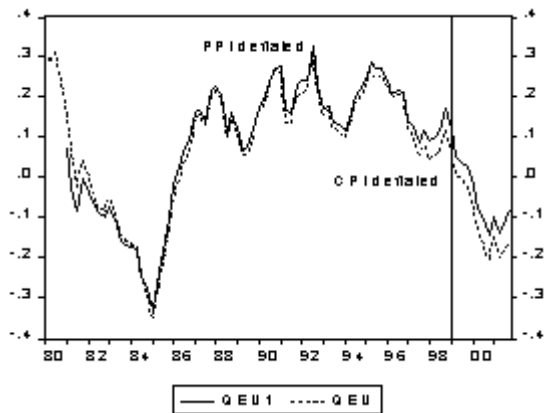


Figure 2. Real Measures of the Dollar/Euro Exchange Rate

exchange rate are largely the same.²

While we have data extending back to 1980/81 for Euro Area GDP and productivity, we decided to focus on the post-1984 sample, largely because none of our candidate variables is able to explain the rise and fall in the dollar during the middle of the 1980's. Thus, we do not have anything to add to the debate over the mid-1980's mystery of “the dazzling dollar”.³

Our first observation is that there is no apparent trend in the real exchange rate (Table 1). Of course, the real exchange rate fails to reject a unit root test, so even if an apparently significant correlation were obtained, it would not necessarily be meaningful.

As mentioned earlier, GDP growth is one factor that has often been viewed as important. Using GDP data extending back to 1985q1, one finds that there is an apparently substantial elasticity of the real rate with respect to the GDP differential of approximately -1.7. This finding appears to validate the conjecture that income differentials are key to the euro weakness. In addition, higher anticipated asset returns in the U.S. are thought to be correlated with this gap. The theoretical link between profitability and the exchange rate is made explicit by Bailey, Millard, and Wells (2001), who argue that an increase in productivity raises future expected profits, raising equity prices and stimulating investment. This additional investment can be financed by capital inflows, enabling domestic residents to finance the additional investment demand without forgoing current consumption. These effects are consistent with other aspects of

² Although as noted in Chinn and Alquist (2000), the PPI deflated dollar/euro rate over the more recent period has exhibited more evidence of stationarity.

³ The title of Frankel's (1985) article on the dollar. The correlation between the real rate and productivity differentials that we identify holds up in regressions over a sample extending back to 1980q1, although the cointegration test results are ambiguous.

the “New Economy” explanation of the appreciation of the dollar with respect to the euro (Meredith, 2001). This explanation argues that high returns on U.S. investments have attracted large foreign inflows of capital, which in turn have appreciated the dollar through capital account effects.

In the next column, we examine a perennial favorite for proxying productivity effects in the Balassa-Samuelson framework: the relative nontraded-traded price differential. An apparently strong relation is detected. Each one percentage point increase in the intercountry nontraded-traded price differential results in a 5.3 percent real appreciation.

Economy-wide productivity differentials also appear to have a strong impact on the dollar/euro rate. In this case, the elasticity is -4.4, that is, a one percentage point increase in the U.S.-euro area productivity differential results in a 4.4 percent appreciation.

In the final column, we present the results of an OLS regression with government spending to GDP ratio as the independent variable. There is apparently only a weak relationship between government spending and the real exchange rate. Assuming government spending falls predominantly on non-tradables, the government spending to GDP ratio possesses the correct sign. As the paper subsequently documents, other procedures couched in terms of a long-run relationship fail to detect a strong correlation between the real exchange rate and government spending.

The correlation between the real exchange rate and the two productivity measure provides the puzzle that this paper addresses: why does productivity matter so much? In most productivity-based models, productivity increases yield appreciations of a magnitude comparable to the share of nontradables in the aggregate basket. Literally, this means that the coefficient on

productivity should be bounded in absolute value between zero and one (see the discussion in Chinn, 1997), which makes it difficult to appeal to measurement error in explaining this finding. Even assuming that both the relative price variable and the broad productivity indices mismeasure the relevant tradable/nontradable productivity differential, the observed elasticity is still more than four times the expected magnitude.⁴ The situation is even worse when one considers competing productivity based models. In the Lucas (1982) and Stockman (1980) models, productivity increases reduce the relative price of home goods, and hence under the simplest assumptions regarding preferences, lead to real depreciations of the currency. Similarly, models in the Ricardian framework (Dornbusch, Fischer, and Samuelson, 1977) suggest that productivity increases lead to depreciations in the home currency.

To see why this is necessarily true, consider a world with an infinite number of differentiated goods. The set of goods produced at home, versus abroad, is defined by relative labor productivity levels; the relative wage rate is endogenously determined in conjunction with the relative labor demand curve. The solution is straightforward when utility is Cobb-Douglas and consumption of goods is proportional to income, which ensures that trade is balanced.

In Figure 3, the A schedule depicts relative unit labor costs across a continuum of goods, ordered from highest to lowest. The B schedule shows demand across these goods. Equilibrium as depicted in Figure 3 is determined by the intersection of the two schedules. 0 to z_l products are produced at home, while z_l to z_m are produced in the foreign country. The relative wage rate is

⁴ Consider what happens if $a = \alpha a^N + (1-\alpha) a^T$, where N (T) denotes nontradable (tradable) sectors, and a is productivity. Then for $\alpha = 0.5$, and with the correlation between a and a^T selected to maximize the estimated coefficient, the probability limit of the regression coefficient on aggregate productivity in a regression of the real exchange rate on productivity is 1.

given by w/w^* . This relative wage rate is also the real exchange rate, given the model's Ricardian structure.

An increase in productivity across the board would be represented by an upward shift of the entire A schedule, resulting in an increase in home relative wages. Since the $B(z)$ curve slopes upward, the increase in equilibrium relative wages is smaller than the effect of the productivity increase.⁵ This means the price of home goods falls and implies a real depreciation, an outcome that runs counter to what has been observed.

If, on the other hand, the productivity growth occurs only in the U.S. there is a kinked shift up of the A schedule to A'. Income increases without changing relative wages and the real exchange rate. One could interpret the advent of the New Economy in this way. As documented by Gust and Marquez (2000), the pickup in labor productivity appears to be largely restricted to the U.S., and has yet to diffuse to the rest of the OECD countries.

⁵ If the $B(z)$ curve is vertical, the increase in equilibrium relative wages exactly offsets the change in productivity.

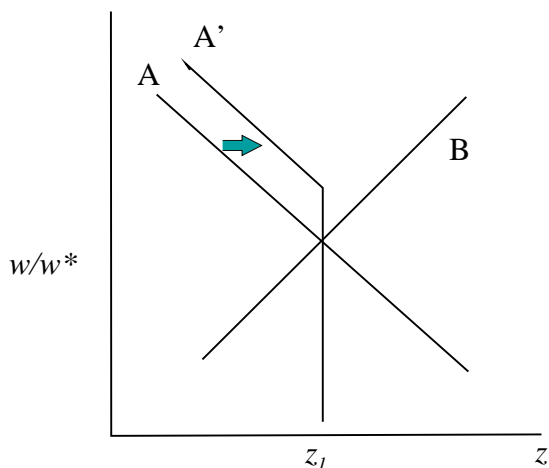


Figure 3. Equilibrium in the Ricardian Model

This leaves the *appreciation*, not merely the strength, of the dollar over the past half decade unexplained. The dollar has been rising on a trade weighted basis and against the euro since early 1995. Tille, Stoffels, and Gorbachev (2001) provide an empirical analysis of the effect of productivity on the euro-dollar rate over three decades using annual OECD data. They find that the Balassa-Samuelson effect accounts for an approximately 1% per year depreciation of the real exchange rate for the 1990s, or about two-thirds of the 1.8% annual depreciation of the real euro-dollar rate over the decade. Clearly the magnitude of the effect of productivity on the real exchange rate documented above is much larger than theoretical considerations warrant. As Meredith (2001) points out, taking the results of Tille, Stoffels, and Gorbachev at face value, one would expect a cumulative real appreciation of the dollar of around 5% from 1995 onwards. Instead the euro lost about 42% of its real value between 1995 and the end of the decade.

To reconcile the euro's depreciation with the Dornbusch-Fischer-Samuelson model, one has to drop the assumption that preferences are homothetic. Instead, one needs to argue that there has been a shift in demand toward U.S. goods, interpreted in the diagram as a shift downward in the *B* schedule. This is a more difficult case to sustain, as the late 1990's have been marked by large trade deficits. However, these could be rationalized to some extent by wealth effects omitted from the Ricardian model.

As any reader of the current literature on the New Economy and the strength of the dollar knows, none of these ideas is new in themselves. Rather, we believe it is useful to categorize some of the many – not necessarily internally consistent – ideas in a framework that is recognizable to economists. We now move to documenting formally the correlation between productivity and the real exchange rate.

3 ECONOMETRIC METHODOLOGY

Two approaches are used in estimating long-run relationships. The first is the Johansen (1988) and Johansen and Juselius (1990) maximum likelihood approach which allows for multiple cointegrating vectors. The other approach is the dynamic ordinary least squares (DOLS) approach of Stock and Watson (1993). The two sets of estimates are used to check for the robustness of the results. On the basis of monte carlo simulation experiments, Stock and Watson report that although the maximum likelihood estimates obtained from the Johansen technique are unbiased, the estimates are also more dispersed than those provided by alternative estimators, such as those obtained from DOLS. In particular, although the DOLS estimates are biased, they have a smaller mean squared error.

3.1 The Johansen Maximum Likelihood Approach

Let x_t be a $m \times 1$ vector of $I(1)$ variables. Then one can estimate the VAR(p):

$$\Delta x_t = \mu + \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \dots + \Gamma_{p-1} \Delta x_{t-p+1} + \Pi x_{t-p} + Bz_t + u_t \quad (1)$$

where $\Gamma_1, \Gamma_2, \dots, \Gamma_{p-1}$, and Π are $m \times m$ matrices of unknown parameters, z_t is a $s \times 1$ vector of $I(0)$ variables (such as time trends) and u is distributed $N(0, \Sigma)$. The matrix Π is estimated by the Johansen maximum likelihood procedure subject to the hypothesis that Π has reduced rank (i.e., $r < m$). This hypothesis is written:

$$H(r): \Pi = \alpha\beta' ; \text{Rank}(\Pi) = r \quad (2)$$

where α and β are $m \times r$ matrices. If $r < m$ then under certain conditions the $\beta'x_t$ is stationary (i.e., the elements of x are cointegrated).

There are two tests proposed by Johansen, and described in greater detail in Johansen and Juselius (1990). The trace statistic for testing $H(r)$ against $H(m)$ is given by:

$$\begin{aligned} J_T &= 2[l_U(\hat{\Theta}) - l_R(\tilde{\Theta})] \\ &= -N \sum_{i=r+1}^m \log(1 - \hat{\lambda}_i) \end{aligned} \quad (3)$$

where Θ is the vector of parameters, and the λ 's are the solution of a determinantal equation involving residuals from a regression of x on first difference terms.

The maximal eigenvalue statistic for testing $H(r-1)$ against $H(r)$ is given by

$$\lambda_{\max} = -N \log(1 - \hat{\lambda}_i) \quad (4)$$

The asymptotic critical values for the case where trends are allowed in the cointegrating vector but not in the data generating process are drawn from Osterwald-Lenum (1992). Cheung and Lai (1993) have shown that finite sample critical values may be more appropriate given the relatively small samples which are generally under study.⁶ Results for both sets of critical values are reported below.

3.2 The Dynamic OLS Approach

Stock and Watson propose a OLS-based estimator for cointegrating relations, which has desirable properties under the conditions that (i) the system is triangular and (ii) there is only one cointegrating relationship.

Let x_t represent a m -dimensional time series, the elements of which are $I(1)$. Suppose $E(\Delta x_t) = 0$, and the $m \times r$ matrix of r cointegrating vectors is $\alpha = (-\Theta, I_r)'$, where Θ is the $r \times (m-r)$ submatrix of unknown parameters to be estimated. The triangular representation for x_t is:

$$\begin{aligned} \Delta x_t^1 &= v_t^1 \\ x_t^2 &= \mu + \theta x_t^1 + v_t^2 \end{aligned} \tag{5}$$

where x_t is partitioned as (x_t^1, x_t^2) , x_t^1 is $(m-r) \times 1$ and x_t^2 is $r \times 1$, and $v_t = (v_t^1, v_t^2)'$ is a stationary stochastic process.

The dynamic OLS estimator is obtained by running the regression:

⁶ The finite sample critical values are obtained by adjusting the asymptotic critical values for the loss of degrees of freedom due to the estimation of the parameters describing the short term dynamics. The adjustment factor is given by $(N - (p \times m))/N > 1$, where N is the number of observations. More directly, the finite sample critical values can be calculated from their equation 11 and the coefficient estimates in Table 1. See Cheung and Lai (1993) for details.

$$\begin{aligned}
x_t^2 &= \beta x_t^1 + \xi_i \sum_{i=-2}^{+2} \Delta x_{t+i}^1 + w_t^2 \\
w_t^2 &= v_t^2 - E[v_t^2 | v_t^1]
\end{aligned}
\tag{6}$$

The leads and lags of the difference terms serve to orthogonalize the error term associated with the dependent variable.

The requirement of recursiveness will not be particularly onerous in the current context since the right hand side variables (such as GDP and labor productivity) can plausibly be interpreted as exogenous.

4 EMPIRICAL RESULTS

4.1 Data

The data are quarterly in frequency, for the period 1985q1-2001q4. The real exchange rate is calculated using consumer price indices. GDP figures are in real constant 1995 units for the euro area and chain weighted 1996 units for the United States. The government spending variable is the log of real government spending minus log real GDP. (Greater detail on the data sources and variable construction can be found in the Data Appendix.). The predicted negative sign on government spending is conditional on most government spending falling upon domestic goods.

Quarterly data on Euro area economy-wide labor productivity levels are calculated from the ECB growth rates for GDP per worker reported in the ECB *Monthly Bulletin*, while those

from the U.S. are the standard non-farm business sector series reported by the BEA.⁷ Oil prices are measured by the world price in dollars deflated by the U.S. CPI from *International Financial Statistics*.

The two measures of productivity differ across the regions. The Euro Area productivity variable is a per worker series which includes all of GDP (agriculture, public sector), while the U.S. series pertains largely to the private sector, and is calculated on a per hour basis. This difference may be problematic if there are trend differences in the behavior of GDP/worker and private sector/hour series. The two series will also tend to diverge during cyclical peaks and troughs.

The problems associated with using different productivity measures are illustrated in an OECD study (Scarpetta, et al., 2000) which demonstrates that one's conclusion about U.S. productivity growth in the 1990s is potentially sensitive to the measure used to assess U.S. performance. In particular, the performance of U.S. productivity growth looks less striking when compared with Europe's if GDP per hour worked is used as opposed to GDP per employee. However, the authors of this study nevertheless partially attribute U.S. growth performance to the productivity improvements in the information and communication technology sector and suggest that the diffusion of these technologies to other industries contributed to the strong output and productivity growth in the U.S. during the 1990s. In any event, the OECD study employs annual productivity data, whereas ours is unique in that it uses quarterly data. To the extent that we are mainly concerned with long-run relationships, the problems associated with using the different

⁷ It might be argued that average labor productivity is tainted by demand effects (labor hoarding, etc.) However, the virtue of couching the analysis in terms of cointegration is that one can sidestep such issues.

measures may be mitigated. Until comparable quarterly productivity data for the U.S. and the Euro Area become available, we are constrained by the data at hand.

Unit root tests were performed on all the relevant variables. The results of these tests are reported in Appendix Table A1. No series rejected the unit root null using ADF tests at the 5% level, except for the real government spending differential and real oil prices. Hence, we proceed assuming the relevant series are $I(1)$. As indicated below, the estimated system that includes the real government spending differential implies the relevant variables are stationary, although the large effect of the productivity differential on the real exchange rate still stands out.

4.2 Cointegration Results

The maximum likelihood results are presented in Table 2. Using the 5% asymptotic critical values, there appears to be evidence for at least one cointegrating vector in the system of real exchange rates and (log) productivity differentials. These conclusions are unchanged if a time trend in the data generating process is allowed for. The elasticity of the real exchange rate with respect to the differential is -5.1, regardless of whether a time trend is assumed in the cointegrating vector or not. Inclusion of oil prices (columns 3 and 4) yields similar though slightly lower estimates of this elasticity, although oil prices themselves do not appear to be an important determinant of the real exchange rate and the Johansen procedure fails to identify a cointegrating relationship between the variables when the finite sample critical values are used.

We compare the results using (log) GDP differentials in place of productivity (columns 5 and 6). The Johansen procedure indicates an absence of cointegration, although higher U.S. income is associated overall with a stonger dollar, which conforms to the conventional wisdom about the

recent behavior of the euro. Hence, it appears that the long run relationship between output and the real exchange rate is not robust, although a short run relation in growth rates might still hold.

Finally, we augment the productivity equation with government spending (columns 7 and 8) and find that although government spending has the expected negative sign, it is not statistically significant and the Johansen procedure fails to detect a long-run relationship between the variables. The effect of productivity on the real exchange rate remains large, however.

As Stock and Watson have pointed out, the DOLS estimator exhibits less dispersion than the Johansen estimator. Hence, we also examine the estimates obtained using their procedure. The DOLS estimates for eight specifications are presented in Table 3. In contrast to some other studies (Chinn, 1997), the productivity coefficient estimates are similar to those obtained using the Johansen procedure: -4.4 to -4.8 (in columns 1-4). The price of oil enters with the expected sign, although it is significant only in the specification including a time trend.

In columns 5 and 6, the productivity differential is replaced with the GDP differential. One finds that the estimated relationship is insensitive to the inclusion of a time trend, suggesting that the correlation in levels is not spurious. This is a surprising pattern of results, given the failure to detect cointegration for the exchange rate-GDP pair using the Johansen procedure.

In columns 7 and 8, we present the results of the productivity equation with government spending included. We find evidence of three cointegrating vectors, even when using finite sample critical values. One interpretation is that the Johansen procedure is detecting the stationarity of the government spending differential and real oil prices, hinted by the univariate tests. Perhaps this is why we fail to identify a close relationship between the real exchange rate and government spending. The the effect of the productivity differential remains large (approximately -2 to -3) and

statistically significant.

It is of interest to compare these results with those obtained by some other recent studies. In particular, the very strong effect of productivity on output is rather remarkable, although not unheard of. Applying the Johansen procedure to annual data from 1960 to 1991, Strauss (1996) estimated tradable/nontradable productivity elasticities of between -1.2 to -8.7. However, these results pertain to OECD countries in general, rather than the constituent currencies of the euro, or the synthetic euro.

In analyses of the synthetic euro, Makrydakakis *et al.* (2000) identified an elasticity of between -0.95 and -1.21 over the 1980-1999 period, where productivity is measured by aggregate GDP divided by employment. Alberola, Cervero, Lopez and Ubide (1999) identifies a smaller (in absolute value) elasticity of -1.64, as does Koen, Boone, de Serres and Fuchs (2001) for the real effective euro, but not for the dollar/euro exchange rate.⁸ In more recent work, Maeso-Fernandez, Osbat and Schnatz (2001), obtain an elasticity of -5.1 in a specification similar to ours.

4.3 Implied Equilibrium Values

The parameter estimates obtained from the constrained specification are economically quite large. They indicate that a 1% increase in US productivity induces approximately a 5% real appreciation in the dollar relative to the euro. Using this parameter⁹, an estimate of the long-run real

⁸ The somewhat lower impact may be attributed to the fact that the productivity variable is measured as GDP/worker, a variable more likely to be comove with GDP than with labor productivity as measured by output per hour. The GDP/employee elasticities are very similar to our GDP elasticities of -1.7 to -3.1 in columns 5 and 6 of Table 3.

⁹ The actual specification conforms to that reported in Table 3, column 3. This assumes a fairly small oil price elasticity of -0.05.

exchange rate is generated. This estimate is illustrated in Figure 4.

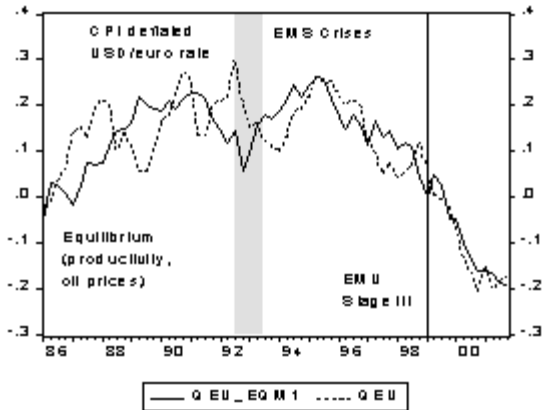


Figure 4: Actual and Equilibrium Real Dollar/Euro Rate

Interestingly, this calculation suggests that the current value of the euro is slightly above its equilibrium value. As of 2001q1, our model implies an equilibrium exchange rate of 0.85 \$/euro, while the actual average value was 0.92 \$/euro, a gap of 8%. This estimated disequilibrium is larger than that identified in other studies, but the estimated gaps will necessarily differ according to their reference period, and the underlying economic model.

There are few other studies estimating the equilibrium in 2001q1. Fels and Yilmaz (2001) use a behavioral equilibrium exchange rate (BEER) model incorporating productivity and real interest differentials and the real price of oil. They conclude that the euro was approximately at the equilibrium value of 0.95 dollars/euro at the beginning of 2001.

One potential criticism of our approach is that the estimate of the dollar/euro equilibrium value may merely be incorporating a fragile correlation between the dollar/euro rate and productivity differentials that has become increasingly pronounced in the past few years, and that might dissipate in the near term. While it is not possible to determine ahead of time whether this relationship will

hold in the future as strongly as in the recent past, it is possible to determine whether using pre-1998 data would have led to drastically different results. Using the 1986-98 sample, one would have obtained a 15% misalignment, instead of the 8% indicated using the entire sample.¹⁰ This discrepancy arises because of the smaller elasticity (-2.7 versus -5.0) and the effect of oil prices.¹¹ Hence, the productivity trends of 1999 and 2000 are not driving our estimates.

6. CONCLUSIONS

This paper has documented the evidence in support of a stable long-run relationship between the real dollar/euro exchange rate, productivity differentials, and the real price of oil. Productivity differentials enter into the relationship in a robust, although unexpected, manner. Productivity has too large an effect to be rationalized within the Balassa-Samuelson framework, and it possesses the incorrect sign for simple versions of two-good neoclassical or Ricardian models. Thus, one needs to incorporate additional assumptions. We believe the key required assumption is that spending falls disproportionately on domestic goods – an assumption commonplace in ad hoc macroeconomic models, but often omitted in neoclassical models for the sake of tractability. Higher income, and expectations of higher income, induce a disproportionate increase in spending on domestic goods, thereby causing an appreciation of the home currency (the dollar in this case) when a depreciation

¹⁰ This specification incorporates elasticities of -2.7 for the productivity differential and -0.03 for oil prices. As one shortens the sample, the productivity elasticity declines in absolute value, and the implied misalignment increases.

¹¹ Several studies refer to the end of 1999. Clostermann and Schnatz (2000) estimate a misalignment of approximately 8%, while Chinn and Alquist (2000) find a 15% misalignment, based on historical correlations from a *monetary* model. Using our specification with productivity and oil prices for a sample extending up to 2000q4, the euro was some 4% undervalued at the end of 1999/beginning of 2000.

would be predicted on the basis of only supply side effects.

In terms of prospects for the dollar/euro rate, the productivity based model that we use is best thought of as a means of assessing the direction of the dollar/euro rate conditional upon a given trend in US-euro area productivity differential. Unfortunately, there is considerable uncertainty about the future path of this trend. Even before the events of September 11th, it was apparent that the Euro area growth was unable to transcend the limits encountered in previous expansions, although there is no clear reason why “New Economy” opportunities should remain confined to the United States. Thus, while the approximately 1.5 percentage point productivity differential in favor of the U.S. observed over the 1995-2000 period is unlikely to be sustained, it is also plausible that a positive productivity differential might persist even as the U.S. economy enters into a recession.¹² If this is the case, the long run value of the euro may continue to decline, although cyclical and portfolio balance factors associated with recent events may obscure that trend.

¹² For arguments that accelerated productivity growth will continue, see Baily (2002) and CEA (2002), Box 1-4 “Is there still a New Economy?”.

DATA APPENDIX

Except otherwise noted, the data are seasonally adjusted quarterly data, derived from the IMF's *International Financial Statistics (IFS)*.

Exchange Rates

Description: quarterly average of monthly bilateral dollar/euro rates.

Income

Description: GDP in 1996 chain weighted dollars (United States) or 1995 constant euros (euro area) for EU-11 before 2000Q4 and for the EU-12 from 2001Q1.

Consumer Price Index

Description: CPI-All items for the US, HICP for EU-11 before 2000Q4 and for the EU-12 from 2001Q1.

Real Exchange Rate

Series: CPI adjusted exchange rate

Description: $\log(\text{spot rate}) - \log(\text{US CPI}) + \log(\text{foreign CPI})$

Productivity

Description: Non-farm business labor productivity for the U.S. Economy-wide labor productivity for EU-11, measured by GDP per employee.

Source: BEA for U.S. and authors' calculations using ECB data for EU-11. For series plotted in Figure 1, data provided by Bernd Schnatz.

Government Spending

Series: Government spending to GDP ratio.

Description: $\log(\text{real government consumption and investment}) - \log(\text{real GDP for U.S.})$, $\log(\text{real government consumption}) - \log(\text{real GDP for EU-11})$.

Source: BEA for U.S., ECB data for EU-11.

Oil Price

Series: Real price of oil

Description: World price of oil.

Source: *International Financial Statistics* line 00176aazzf...

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TABLE 1
Summary Correlations
Dependent Variable: Real Exchange Rate

	(1)	(2)	(3)	(4)	(5)
Time	0.04*** (0.00)				
Time ²	-0.00*** (0.00)				
GDP		-1.73*** (0.54)			
Prod'y			-5.26*** (0.79)	-4.41*** (0.40)	-4.57*** (0.54)
Gov't					-0.18* (0.09)
\bar{R}^2	.77	.21	.40	.49	.51
N	68	68	68	68	68
Sample	1985q1- 2001q4	1985q1- 2001q4	1985q1- 2001q4	1985q1- 2001q4	1985q1- 2001q4

S))
 Notes: Coefficient is from OLS regression over the indicated sample. Productivity in column 3 is proxied by the ratio of the CPI to PPI. Productivity in columns 4 and 5 is measured by the ratio of US nonfarm business labor productivity to Euro area economywide labor productivity.

TABLE 2
Johansen Cointegration Results for Neoclassical Model

	(1)	(2)	(3)	(4)
lag	5	5	1	1
LR	23.0	22.7	36.5	35.2
c.v.	1[1]	1[1]	0[1]	1[1]
Coefficient				
prod.	-5.14 (0.66)	-5.14 (0.67)	-4.67 (0.87)	-4.68 (0.88)
oil price			-0.01 (0.08)	-0.01 (0.08)
time trend	no	in data	no	in data
N	63	63	66	66
sample	1986q2- 2001q4	1986q2- 2001q4	1985q3- 2001q4	1985q3- 2001q4

Notes: Entries under LR indicated the likelihood ratio test statistic for the Maximal Eigenvalue test; c.v. indicate the number of cointegrating vectors indicated using the 5% critical values adjusted for sample size as per Cheung and Lai (1993). The numbers in brackets [.] are the implied number of vectors using the asymptotic critical values from Osterwald-Lenum (1990). Coefficients are the parameters from the estimated cointegrating vector, normalized on the real exchange rate. If zero vectors are indicated, the estimated cointegrating vector is calculated assuming one cointegrating vector. Lag length is determined by the Akaike Information Criterion (AIC). The AIC picks one lag in levels and thus no lags in the VECM for the regressions in columns 3, 4, 7, and 8. In these cases, we augment the VECM with one lag in first differences.

TABLE 2 (continued)
Johansen Cointegration Results for Neoclassical Model

	(5)	(6)	(7)	(8)
lag	2	2	1	1
LR	14.7	13.7	87.0	74.3
c.v.	0[0]	0[0]	3[3]	3[3]
 Coefficient				
GDP	-3.31 (1.07)	-3.29 (1.07)		
prod.			-2.26 (0.82)	-2.91 (0.71)
gov't.			-0.25 (0.17)	-0.09 (0.25)
oil price			-0.23 (0.08)	-0.20 (0.07)
time trend	no	in data	no	in data
N	68	68	66	66
sample	1985q1- 2001q4	1985q1- 2001q4	1985q3- 2001q4	1985q3- 2001q4

Notes: Entries under LR indicated the likelihood ratio test for the Maximal Eigenvalue test; c.v. indicate the number of cointegrating vectors indicated using the 5% critical values adjusted for sample size as per Cheung and Lai (1993). The numbers in brackets [.] are the implied number of vectors using the asymptotic critical values from Osterwald-Lenum (1990). Coefficients are the parameters from the estimated cointegrating vector, normalized on the real exchange rate. If zero vectors are indicated, the estimated cointegrating vector is calculated assuming one cointegrating vector. Lag length is determined by the Akaike Information Criterion (AIC). The AIC picks one lag in levels and thus no lags in the VECM for the regressions in columns 3, 4, 7, and 8. In these cases, we augment the VECM with one lag in first differences.

TABLE 3
Stock-Watson Cointegration Results for Neoclassical Model

	(1)	(2)	(3)	(4)
Coefficient				
prod.	-4.82*** (0.65)	-4.79*** (0.70)	-4.85*** (0.66)	-4.45*** (0.58)
oil price			-0.05 (0.05)	-0.18*** (0.05)
time trend	no	yes	no	yes
adj.-R ²	.67	.66	.65	.68
N	63	63	63	63
sample	1985q4- 2001q2	1985q4- 2001q2	1985q4- 2001q2	1985q4- 2001q2

Notes: Coefficient indicates the estimated DOLS coefficient, where the estimator corresponds to DOLS2 in Stock and Watson (1993). Newey-West standard errors using the Bartlett window and first four lags. *(**)(***) denotes significance at the 10%(5%)(1%) level.

TABLE 3 (continued)
Stock-Watson Cointegration Results for Neoclassical Model

	(5)	(6)	(7)	(8)
Coefficient				
GDP	-1.69*** (0.59)	-3.05*** (0.83)		
prod.			-4.23*** (0.47)	-3.90*** (0.98)
gov't.			0.08 (0.25)	-0.43 (0.61)
oil price	-0.05 (0.05)	-0.18*** (0.05)	-0.14** (0.07)	-0.13 (0.08)
time trend	no	yes	no	yes
adj.-R ²	.13	.26	.89	.89
N	66	66	42	42
sample	1985q1- 2001q2	1985q1- 2001q2	1991q1- 2001q2	1991q1- 2001q2

Notes: Coefficient indicates the estimated DOLS coefficient, where the estimator corresponds to DOLS2 in Stock and Watson (1993). Newey-West standard errors using the Bartlett window and first four lags. *(**)(***) denotes significance at the 10%(5%)(1%) level.

TABLE A1
Unit Root Tests
1985.1-2001.4

variable	τ_c	τ_τ
q	-2.66	-2.34
a ^{US}	1.13	-1.12
a ^{EU}	-2.06	-1.25
a ^{US} - a ^{EU}	-0.90	-1.35
gdp ^{US}	-1.38	-2.40
gdp ^{EU}	-1.56	-1.99
gdp ^{US} - gdp ^{EU}	-1.31	-2.34
gov ^{US}	-1.77	1.92
gov ^{EU}	-2.04	-3.51
gov ^{US} - gov ^{EU}	-4.44*	-1.50
p ^{oil}	-3.00*	-3.25

Notes: τ_c (τ_τ) is the ADF statistic with constant (constant and trend), including 4 lags of first differences. * indicates significance at the 5% MSL.