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Inflation, Money Demand, and Purchasing Power Parity in South Africa

GUNNAR JONSSON*

This empirical study for South Africa indicates that there exists a stable money demand type of relationship among domestic prices, broad money, real income, and interest rates, as well as a long-run relationship among domestic prices, foreign prices, and the nominal exchange rate. In the short run, shocks to the nominal exchange rate affect domestic prices but have virtually no impact on real output, while shocks to broad money have a temporary impact on real output before becoming inflationary. Both types of shocks seem to trigger a monetary policy response, as the short-term interest rate adjusts quickly. [JEL C32, E31, E41, F41]

South Africa adopted a formal inflation-targeting framework for monetary policy early in 2000, following less than satisfactory experiences with other monetary policy regimes (such as an exchange rate peg and money growth targeting, see Box 1) during the previous decades. The inflation target was set at 3–6 percent by 2002, and transparency and accountability of the South African Reserve Bank (SARB) were enhanced.

The introduction of the new monetary policy regime was done in the context of a considerable fall in inflation during the 1990s. Annual growth in the underlying consumer price index fell from 18 percent in 1991 to 13 percent in 1993 and further to 7 percent in 1998. Developments in broad money (M3), however, did

*Gunnar Jonsson is a Senior Economist at the International Monetary Fund. He wishes to thank Trevor Alleyne, Torbjörn Becker, David Coe, Neil Ericsson, José Fajgenbaum, Zenon Kontolemis, Michael Nowak, Jean-Claude Nachega, Eric Schalling, Arvind Subramanian, Yougesh Khatri, Krishna Srinivasan, Carl Walsh, Tarik Yousef, and seminar participants at IMF's African Department and the University of Cape Town for discussions and comments. The usual disclaimer applies.

Box 1. Monetary and Exchange Rate Regimes in South Africa

The South African Reserve Bank has operated different monetary and exchange rate policy regimes during 1970–2000. This period has also witnessed a substantial degree of financial and external liberalization, where the latter included both trade reforms and capital control liberalization.

1970–79: Between 1970–79, the rand was pegged to either the U.S. dollar or the pound sterling. However, frequent alterations to the level of the peg were undertaken in the form of discrete step changes. The Reserve Bank also used changes in cash and liquid asset requirements combined with credit ceilings and interest rate controls to affect liquidity conditions. At the same time, exchange controls severely restricted the capital flows of residents, while nonresidents had to place the proceeds from sales of South African assets in blocked rand accounts, which could only be freely transferred overseas after five years.

1980–85: A shift in the monetary policy regime took place in the early 1980s; the Reserve Bank moved to a system of indirect control of money supply and adjustments of short-term interest rates in an effort to enhance the responsiveness of monetary aggregates to macroeconomic developments. Greater flexibility was also introduced into the foreign exchange and capital markets. A managed float but dual exchange rate system was in place between 1979–83, with financial transactions by nonresidents being valued at a discounted exchange rate (the “financial rand mechanism”), while current account transactions were valued at the commercial exchange rate. The liberalization process continued between 1983–85, including the adoption of a unified exchange rate system, scaling down of liquid asset requirements, and dismantling of interest rate controls.

1985–94: In the context of the political upheavals in the mid-1980s, however, South Africa declared a moratorium on most of its debt obligations in 1985, following the refusal by a number of international banks to roll over short-term loans to South Africa. The ensuing financial sanctions resulted in a debt standstill and, subsequently, in a series of rescheduling agreements between 1985–94. Moreover, the financial rand mechanism was reintroduced, and capital controls were effectively tightened. Starting in 1986, the Reserve Bank announced target ranges for growth in broad money, which was lowered over time in an attempt to bring inflation down. Indeed, broad money growth and inflation fell substantially during this period, although the Reserve Bank often missed the explicit money growth target.

1994–2000: Following the general elections in 1994, the new government intensified the liberalization efforts. The financial rand mechanism was terminated in 1995 and the exchange rate unified; capital controls on residents were gradually liberalized and virtually all controls on nonresidents were removed. Moreover, trade tariffs were lowered and exports and imports grew sharply. At the same time, a substantial degree of financial deepening took place, as low-income households gained access to formal banking services to a larger extent. In this context, growth in broad money accelerated and exceeded the target ranges by wide margins every year between 1994–99, but inflation was contained. The Reserve Bank emphasized that the money growth target should be interpreted as an informal guideline. In practice, a more eclectic approach was followed, which involved monitoring a number of different indicators, including various price indices, the shape of the yield curve, the nominal exchange rate, and the output gap. In February 2000, South Africa adopted a formal inflation-targeting framework for monetary policy.

Sources: Aron, Elbadawi, and Kahn (1997); Garner (1994); Moll (1999b); and the South African Reserve Bank (1998).

not follow the same pattern; although the annual growth rate of M3 fell from 14 percent in 1991 to 5 percent in 1993, it increased to 17 percent in 1998. At the same time, the nominal exchange rate has fluctuated widely in South Africa. For example, the nominal effective exchange rate depreciated on average by 6½ percent a year between 1990 and 1995, but by 22 percent and 19 percent in 1996 and 1998, respectively. The latter events were followed by a pickup in inflation, despite a tightening of monetary policy.

Issues related to the determinants and forecasting of inflation have assumed greater importance under the inflation-targeting framework, including the impact of fluctuations in money and the nominal exchange rate. In particular, the contrasting developments in inflation and money growth during the 1990s led analysts to question whether a stable relationship between these two aggregates exists, and whether money demand is stable. At the same time, it has been noted that movements in foreign prices and the nominal exchange rate are likely to have contributed to inflation developments in South Africa, although the specifics of such a relationship have not been examined thoroughly.[1]

The purpose of this study is to examine empirically the relationship among prices, money, and the exchange rate in South Africa within a simple structural setting. More specifically, the study first examines the long-run stability of two economic relationships involving the above-mentioned variables: a money demand type of relationship and purchasing power parity (PPP). Secondly, the short-run responses and comovements among nominal and real variables following various types of shocks are investigated, with a particular focus on how inflation adjusts to these shocks. In the course of doing this, the issues of a potential structural break in the data since 1994—the starting year of the successful political transformation of the economy and the lifting of sanctions—and whether it is appropriate to focus on a more narrow or broader definition of money when estimating money demand are tentatively examined.[2]

From a methodological perspective, it can be noted that the two long-run relationships mentioned above are estimated simultaneously by using a structural vector error-correction model (VECM). This contrasts with most of the literature on PPP and money demand, where (error-correcting) single equation models are estimated. As both relationships involve domestic prices, however, it seems preferable to model the interaction among the variables within a multivariate cointegration context.[3]

[1]It should be noted that South Africa remained a fairly open economy during the 1970s and 1980s, notwithstanding long periods of international trade and financial sanctions. For example, the sum of merchandise exports and imports remained at about 35 percent of GDP during the sanctions period 1985–95, although the financial sanctions forced South Africa to shift from running external current account deficits in the early 1980s to current account surpluses from 1985 to the early 1990s; see, for example, Jonsson and Subramanian (2000) and Lipton (1998) for discussions.

[2]From a policy perspective, it would be important to also examine how several other variables—such as wages, fiscal variables, and capacity utilization—are related to inflation developments in South Africa. This is, however, beyond the scope of the current paper.

[3]Becker (1999) and Price and Nasim (1999) are two recent studies that use a very similar methodological approach to study the issues of PPP and money demand.

The results indicate that there exists a stable and plausible money demand type of relationship among domestic prices, broad money, real income, and interest rates, as well as a long-run relationship among domestic prices, foreign prices, and the nominal exchange rate. In the short run, it is found that shocks to the exchange rate affect domestic prices but have virtually no impact on real output. This contrasts with shocks to broad money, which have a temporary impact on real output before inflation picks up. Both types of shocks seem to trigger a monetary policy response, as the short-term interest rate adjusts quickly.

I. Background, Methodology, and Data

Empirical Background

The empirical literature on money demand and PPP is substantive, but in most studies the focus is typically on only one of the two relationships.[4] This is also the case for South Africa. The most recent studies on the demand for money in South Africa are Hurn and Muscatelli (1992) and Moll (1999a); both studies find a sensible demand function for broad money despite a degree of financial innovation and liberalization in the 1980s and 1990s.[5] This is corroborated by DeJager and Ehlers (1997) who show that growth in M3 is a better and more stable indicator for future inflation rates than narrow money, and that M3 has a consistent negative relationship with interest rates. These results contrast with Doyle (1996), who tentatively argues that narrow money (notes and coin in circulation outside the banking system) might warrant a more prominent role in the monetary policy framework, as this aggregate is a fair leading indicator for inflation, and as the demand for narrow money appears to be stable.

Issues related to whether PPP holds in South Africa have been examined in various papers. Tsikata (1998) and Subramanian (1998) show that the effective nominal depreciation of the rand during the 1990s is almost fully reflected in higher prices of imported goods, although the results are sensitive to the choice of price aggregates and sample period. This result does not necessarily mean that PPP holds for national price levels. Indeed, Aron, Elbadawi, and Kahn (1997) argue that the real exchange rate in South Africa is nonstationary, implying that a strict interpretation of PPP would not hold. They show, however, that fluctuations in the real exchange rate can be explained by variations in a set of economic "fundamentals," including measures of trade liberalization, terms of trade, government expenditures, capital flows, and official reserves.

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[4]Johansen and Juselius (1990), Hendry and Ericsson (1991), and Ericsson (1998) are examples of useful studies that discuss a range of econometric and time-series issues that arises in studies of money demand. MacDonald (1995), Rogoff (1996), and Habermeier and Mesquita (1999) are examples of studies that survey the PPP literature and provide some new results using cointegration methods.

[5]Hurn and Muscatelli (1992) point out that a number of the earlier empirical studies on money demand in South Africa did not estimate the long-run elasticites in an econometrically satisfactory way.

Theoretical Background

The theoretical underpinning for the study of money demand and PPP is standard. The simplest form of the PPP theory suggests that goods market arbitrage enforces parity in national price levels. Hence, converted to a common currency, national price levels should be equal, that is, all variables expressed in natural logarithms,

$$p = q - e, \tag{1}$$

where p is the domestic price level, q represents (effective) foreign prices, and e is the nominal (effective) exchange rate (defined as foreign currency per domestic currency).[6] At the same time, equilibrium in the money market implies that real money supply ($m^s - p$) equals real money demand, with the latter assumed to a positive function of real income, y, and a negative function of the nominal interest rate, i,[7] that is,

$$m^s - p = m^d(y,i). \tag{2}$$

Thus, we end up with a system of six interrelated variables [p, q, e, m, y, i], where economic theory suggests that two long-run relationships could be found: one between domestic prices, foreign prices, and the nominal exchange rate; and another between domestic prices, money, real income, and the nominal interest rate. While we would expect both the real exchange rate and real money demand to be fairly stable in the long run, we would also expect temporary deviations from these two long-run equilibria to affect future fluctuations in the variables such that the long-run equilibria are restored.

In addition to these considerations, a dummy variable for the period 1994–98 was added to the model in an attempt to identify a possible structural break associated with the economic effects of the political transformation that took place in the early 1990s. This transformation, as well as some important economic structural reforms, could have arguably affected both the long-run money demand relationship and the real exchange rate, since it led to both some financial deepening (as low-income households gained access to formal banking services to a larger extent), as well as a strong increase in foreign competition, which in turn could have had a one-off effect on the domestic price level.

[6]A variant of the simplest PPP hypothesis suggests that expression (1) should only hold for tradeable goods, while prices for nontradables in part depends on relative productivity levels. The current study does not make a distinction between prices of tradables and nontradables.

[7]More precisely, economic theory suggests that demand for money depends on the opportunity cost of holding money. Although the opportunity cost for holding cash is larger when the nominal interest rate is higher, it is ambiguous whether broader definitions of money are positively or negatively related to the nominal interest rate, as broad money typically is interest bearing. See the first part of Section II for further discussion.

Data Issues and Econometric Methodology

The empirical analysis was carried out using quarterly data between 1970:1 and 1998:2. All series are plotted in Figure 1. The underlying consumer price index, *p*, was used rather than the headline consumer price index throughout the study. By using underlying inflation, which excludes highly volatile food prices and housing costs (mainly mortgage costs) from the consumer price index, it was expected that the signal-to-noise ratio would improve in the estimations. To further examine whether broad money or narrow money is more closely related to inflation, two different monetary aggregates were used in the analysis: *nc*, notes and coin in circulation outside the banking sector; and *m*3 (broad money), consisting of *nc* plus check and demand deposits, and medium- and long-term deposits. Two alternative interest rates were included in the model: a short-term rate (*i-short*) and a long-term rate (*i-long*). The nominal exchange rate, *e*, and the foreign price level, *q*, were calculated in effective terms, using the Reserve Bank's weights; an increase in the effective nominal exchange rate means an appreciation of the rand. See Appendix for further data details.

Traditional unit root tests (see Table 1) indicated that all series are integrated of order 1, possibly with the exception of foreign prices, *q*, i.e., the series are nonstationary in levels but stationary in first differences.[8] The nonstationarity of the data together with the notion that none of the variables *a priori* can be regarded as exogenous, suggested that an appropriate methodology would be to start with a non-structural vector auto regression model (VAR), and use cointegration tests to examine whether any long-run relationships exist among the variables.[9] As a second step, economic theory (as described above) was used for identification, turning the empirical model into a structural VAR, and specific cointegrating vectors—related to the PPP and money demand hypothesis—were estimated and tested. Hence, to allow for a dynamic interaction among the variables in the system, the two long-run relationships (as suggested by theory) were estimated separately but simultaneously, while no constraints were placed on the short-run adjustments.

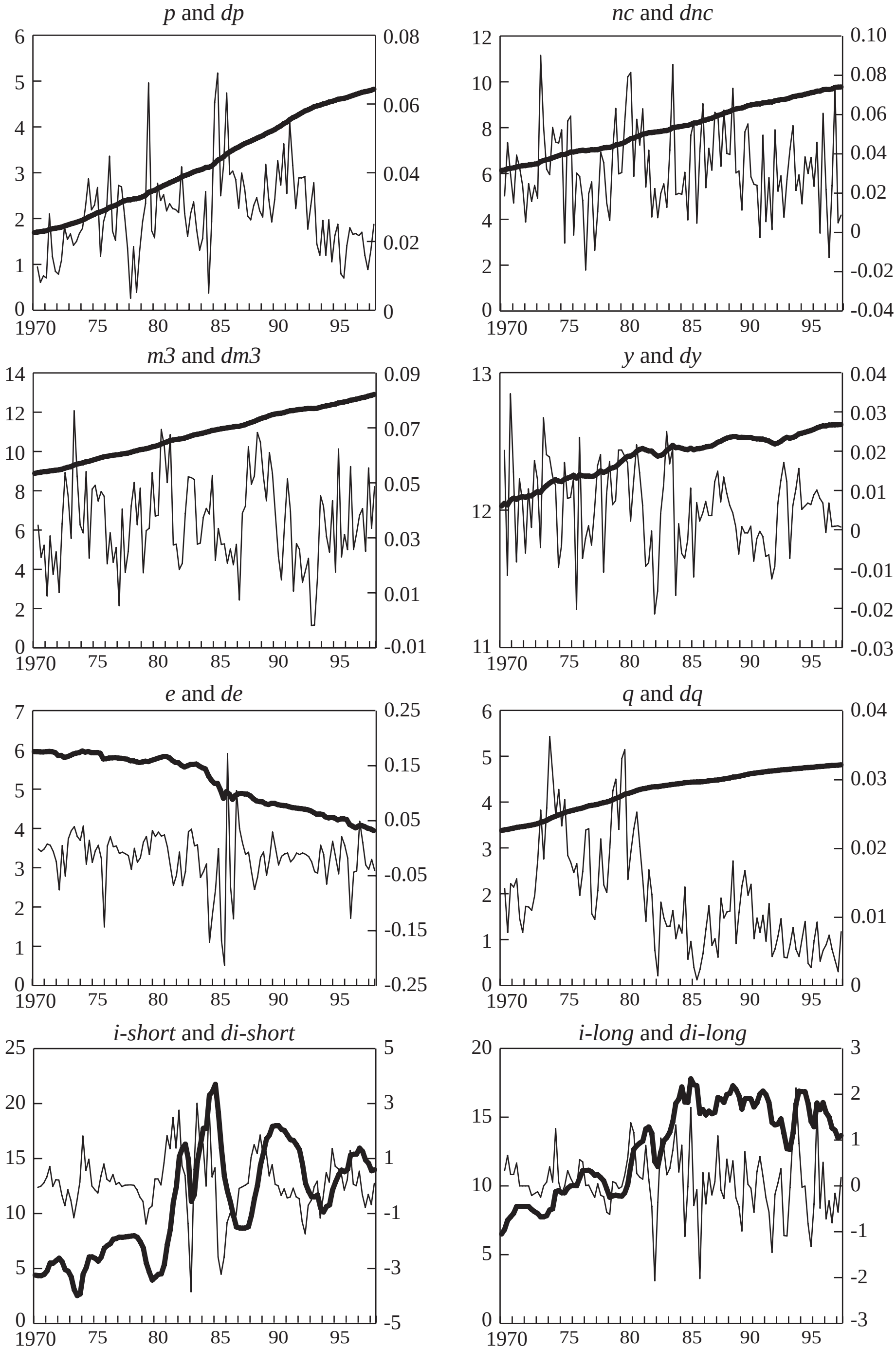
More specifically, following Johansen and Juselius (1990) and Johansen (1991), a vector of endogenous variables, *x*, that are integrated of order 1, is analyzed using the vector error-correction representation,

$$\Delta x_t = \mu + \sum_{i=1}^{k} \Gamma_i \Delta x_{t-i} + \pi x_{t-1} + \varepsilon_t, \tag{3}$$

[8]Visual inspection suggests that a time trend arguably should be included in the first difference test for *q*. Allowing for such a trend also indicates that this series is integrated of order 1. Moreover, the foreign price series adjusted for the effective exchange rate is clearly integrated of order 1.

[9]Foreign prices could arguably be treated as an exogenous variable. Indeed, as shown below, the empirical results reveal that foreign prices do not respond to deviations from the estimated long-run relationships.

Figure 1. Levels and First Differences of the Data Series



Note: Natural logarithms of levels (except for *i-short* and *i-long*), in bold, on left-hand scale; first differences on right-hand scale.

Table 1. Unit Root Tests, 1972:2–1998:2

Variable	*t*-value[1]	Lags Included[2]	Additional Regressors
Levels			
p	–1.67	3	Constant and trend
e	–2.76	3	Constant and trend
q	–1.79	5	Constant and trend
nc	–3.03	4	Constant and trend
*m*3	–2.59	2	Constant and trend
y	–2.43	1	Constant and trend
i-short	–3.24	1	Constant and trend
i-long	–2.21	1	Constant and trend
First differences[3]			
dp	–6.23*	0	Constant
de	–4.29*	2	Constant
dq	–2.01	4	Constant
dnc	–9.27*	0	Constant
*dm*3	–6.86*	0	Constant
dy	–7.46*	0	Constant
di-short	–5.20*	0	Constant
di-long	–7.85*	0	Constant

[1]The *t*-value is the test statistic from the (Augmented) Dickey-Fuller test; * indicates rejection of the null hypothesis of nonstationarity at the 5 percent significance level.

[2]The lag length was chosen by using the Schwarz Bayesian Criterion assuming a maximum of eight lags.

[3]Including a time trend in the first difference regressions did not alter the results qualitatively, except for *q* for which nonstationarity was rejected if a trend was included.

where the parameters μ and $\Gamma_1, \ldots, \Gamma_k$ are allowed to vary without restrictions, k is the lag length of the model, and ε_t is a vector of normally distributed shocks with mean zero. The presence of cointegration is tested by examining the rank of π. In the event of reduced rank of π (that is, when rank $(\pi) = r < n$, where n is the number of endogenous variables), there exists r cointegrating vectors, and the matrix π can be written as $\pi = \alpha\beta'$, with β containing the r cointegrating vectors, and α describing the speed of adjustments to the long-run equilibria (the error-correcting terms). If $r > 1$, the issue of identification arise. In the current paper, the expected rank is 2, implying that (over)identifying restrictions should be placed on the parameters in

$$\pi x_{t-1} = \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \\ \alpha_{41} & \alpha_{42} \\ \alpha_{51} & \alpha_{52} \\ \alpha_{61} & \alpha_{62} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{12} & \beta_{13} & \beta_{14} & \beta_{15} & \beta_{16} \\ \beta_{21} & \beta_{22} & \beta_{23} & \beta_{24} & \beta_{25} & \beta_{26} \end{bmatrix} \begin{bmatrix} p \\ q \\ e \\ m \\ y \\ i \end{bmatrix}_{t-1} . \tag{4}$$

In the estimations below, β_{11} and β_{21} will be normalized to 1, while the simplest forms of the PPP and money demand relationships will be tested by adding exclusion restrictions on [β_{14} β_{15} β_{16}] and [β_{22} β_{23}], respectively.

The short-run comovements among the variables are then examined by generating orthogonalized impulse response functions (see Sims, 1980). As opposed to the traditional VAR literature, however, the computation of the impulse response functions is based on the VECM representation where the estimated long-run restrictions are taken into account. This allows us to examine the effect of a variable-specific shock on the individual variables as well as on the estimated cointegrating relationships; see Pesaran and Shin (1996). The main focus is on shocks to the money, exchange rate, price, and output equations, respectively, in equation (4), with a special emphasis on the inflationary impacts of the different shocks. It should be emphasized that one has to be cautious in interpreting these shocks. For example, a shock to the money equation can originate from a number of different sources, and does not necessarily mean that monetary policy has changed. Likewise, a shock to real output could be due to either an aggregate demand or an aggregate supply shock. See, for example, Becker (1999) for a discussion of these identification issues.

II. Results

Results Regarding the Long-Run Relationships

The results from the first set of cointegration tests are summarized in Table 2. The number of cointegrating vectors was estimated using the Johansen (1988, 1991) procedure.[10] It is well known that cointegration tests in the Johansen setting are sensitive to the lag-length of the VAR. Although it is common to include four lags in the VAR when quarterly data are used, at this stage, the results are reported with two, three, and four lags included in the VAR, respectively.[11] The economic model suggests that two cointegrating vectors should be found (at least as long as only one interest rate is included in the model), and the cointegration tests typically picked up 2–3 stationary vectors (see Table 2, Column 2). The results varied between zero and four vectors, however, depending on the number of lags included in the model, as well as on the choice of monetary aggregate and interest rate. Despite these somewhat inconclusive results, restricted cointegration tests were performed under the assumption of the presence of two cointegrated vectors. The parameters in the restricted model were constrained to test whether the two stationary vectors

[10]The number of cointegrating vectors was estimated using both the maximum eigenvalue statistic and the trace statistic (allowing for unrestricted intercepts but no trends), with the significance level set to 5 percent.

[11]A time dummy for the period 1994:1–1998:2 and seasonal dummies were included in the model as well. The time dummy was included without restricting it to the cointegrating vector, implying that the average growth rates of the variables can change at the time of the structural change, while the cointegrating vectors remain unchanged.

Table 2. Structural VAR-Models, 1971:1–1998:2[1]

Lags	Number of Cointegrating Vectors[2]	Chi-Square Statistic of Likelihood Ratio Tests of:[3] PPP	MD	Joint	Restricted Cointegrating Vectors[4]						
					VAR-models including broad money, *m*3						
					p	*e*	*q*	*m*3	*y*	*i-long*	*i-short*
2	1, 2	1.38	0.34	2.55	1	**0.89**	**–1.26**	0	0	0	
					1	0	0	**–0.76**	**1.14**	**–0.17**	
3	2, 2	2.02	0.61	4.65	1	**0.90**	**–1.26**	0	0	0	
					1	0	0	**–0.93**	**1.82**	**–0.11**	
4	2, 2	0.81	6.83**	7.87*	1	**0.92**	**–1.26**	0	0	0	
					1	0	0	**–0.96**	**1.57**	**–0.10**	
2	3, 3	5.15	3.43	10.34*	1	**0.87**	**–1.27**	0	0		0
					1	0	0	**–1.11**	**1.72**		**0.02**
3	3, 3	5.76	2.84	9.50*	1	**0.91**	**–1.26**	0	0		0
					1	0	0	**–1.08**	**1.43**		**0.03**
4	1, 2	1.86	8.38**	12.28**	1	**0.92**	**–1.12**	0	0		0
					1	0	0	**–1.10**	**1.89**		**0.02**
					VAR-models including narrow money, *nc*						
					p	*e*	*q*	*nc*	*y*	*i-long*	*i-short*
2	2, 3	4.89	1.13	14.04**	1	**0.93**	**–1.23**	0	0	0	
					1	0	0	**–1.18**	**0.68**	**0.15**	
3	2, 2	1.99	2.79	13.36**	1	**0.95**	**–1.24**	0	0	0	
					1	0	0	**0.42**	**3.50**	**–1.46**	
4	1, 1	1.28	5.59*	16.14*	1	**0.96**	**–1.22**	0	0	0	
					1	0	0	**–0.82**	**0.42**	**–0.26**	
2	4, 4	8.18*	5.54*	20.76**	1	**0.87**	**–1.34**	0	0		0
					1	0	0	**–0.98**	**0.07**		**0.01**
3	4, 4	7.07*	0.09	7.59	1	**0.96**	**–1.21**	0	0		0
					1	0	0	**–1.00**	**–0.04**		**0.02**
4	0, 1	2.77	1.49	2.89	1	**0.91**	**–1.23**	0	0		0
					1	0	0	**–0.96**	**–0.08**		**0.01**

[1]The VAR also include (unrestricted) seasonal dummy variables and a time dummy for the period 1994:1–1998:2.

[2]Number of cointegrating vectors is based on Johansen's Trace statistic and maximum eigenvalue statistic, respectively, at the 5 percent significance level.

[3]* and ** indicate rejection of the LR-test at the 5 percent and 1 percent significance level, respectively.

[4]The estimations assume two cointegrating vectors. Bold figures are estimated coefficients.

could be represented by the two long-run relationships discussed above. The results were in part supportive of the theoretical arguments: the hypothesis that one of the cointegrating vectors includes only the variables *p*, *q*, and *e*, was rejected in only two of the twelve specifications (Table 2, Column 3),[12] and the hypothesis that one cointegrating vector includes only the variables *p*, *m*, *y*, and *i* was rejected in four of the twelve specifications (Table 2, Column 4). The joint test of the two hypotheses, however, was rejected in eight of the twelve specifications.

Turning to the parameters of the cointegrating vectors,[13] it can be noted that the estimated parameters for the nominal exchange rate and foreign prices have the expected signs and are fairly close to –1 and 1, implying that the so-called "strict PPP" hypothesis possibly holds. Indeed, it is interesting to note that the joint movements in the nominal effective exchange rate and foreign prices seem to be almost fully reflected in domestic prices in the long run, in the sense that the sum of the estimated parameters (in absolute values) is relatively close to 2. A possible explanation for these results is that domestic price setters sometimes hesitate to adjust domestic prices in line with exchange rate fluctuations—perhaps because they regard these fluctuations as temporary. This would explain a coefficient of less than 1 for *e*. But since this behavior would erode competitiveness in the long run, the price setters compensate by increasing domestic prices by slightly more than a corresponding increase in international prices.

The estimated coefficients in the money demand relationship have the expected signs and are of a plausible magnitude when broad money, *m*3, is included in the model. The estimated coefficient on *m*3 is between 0.8 and 1.1, and the coefficient on real income is between 1.1 and 1.9. The results were less encouraging when narrow money, *nc*, was included in the model. The estimated coefficient on real income was quite unstable and often had the wrong sign, and when the coefficient on *nc* was constrained to equal –1, the estimated income elasticity became even more implausible (not reported). Consequently, a plausible and stable long-run money demand relationship for narrow money could not be established.

The rejection of the joint hypothesis when *m*3 is included in the model could possibly be the result of a misspecification. As broad money to some extent is an interest-bearing asset, it would be preferable to include measures of both the "own" rate of return of this asset as well as the opportunity cost of holding it (see Ericsson, 1998). Hence, the model that included *m*3 was reestimated including both the short-term and long-term interest rates, with the hypothesis that *i-short* would be a proxy for the own rate of return for broad

[12]By not constraining the coefficients on *q* and *e*, the test allows for various fixed costs, such as transportation and menu costs, to vary over time and across countries. The interpretation of this test is simply that the series *p*, *e*, and *q* do not drift too far away from each other. A stricter test of PPP imposes the homogeneity and symmetry restrictions that the coefficients on both *q* and *e* equal 1 (in absolute values), see MacDonald (1995).

[13]The reported parameters in Table 2 are estimated under the assumption of two cointegrating vectors, with exclusion restrictions placed on the β-matrix as discussed in the previous section.

money, while *i-long* would measure the rate of return of alternative assets.[14] A test for system reduction of the number of lags in the VAR indicated that four lags should be included (the results are reported in Jonsson, 1999). The cointegration tests from this model are shown in Table 3. Both the maximum eigenvalue statistic and the trace statistic now indicate that there are three cointegrating vectors in the system. In addition to the two long-run relationships discussed above, it is conceivable that the two interest rates together form another long-run relationship—possibly along with domestic prices and/or the exchange rate; the theory of the term structure of interest rates suggests that a linear combination of long- and short-term interest rates contains information about the future monetary policy stance and could help in extracting expectations about future inflation rates and/or nominal exchange rate movements (see, for example, Mishkin, 1990). Although formal tests of, for example, the Fisher hypothesis or an examination of the term structure of interest rates is beyond the scope of this paper, it is still possible to test the joint hypothesis about the two vectors relating to the PPP and money demand relationships and still allow for the existence of a third cointegrating vector.[15]

The results with regard to the money demand and PPP type of relationships were now supportive of the theoretical arguments. The joint test of the hypothesis that two of the three cointegrating vectors can be represented by a money demand and PPP type of relationship could not be rejected (see Table 3), and the estimated coefficients had the expected sign and were of plausible magnitudes. A visual inspection of the two restricted cointegrating vectors (denoted CV_{ppp} and CV_{md}) further indicates that these vectors seem to be reasonably stationary, see Figure 2 and Table 4.[16]

Regarding the PPP relationship, the unconstrained coefficients on *e* and *q* are estimated to 0.95 and –1.19. Constraining the coefficient on *q* to –1, effectively estimating how the nominal effective exchange rate relates to price differentials in the long run, yielded an estimated coefficient on *e* of 1.06, which is not significantly different from 1. Put differently, the model where the coefficients on *e* and *q* were constrained to 1 and –1, respectively, was not rejected.

The estimated coefficients in the money demand type of relationship were also quite sensible. Long- and short-term interest rates entered the vector with different and predicted signs; that is, higher short-term rates are positively related to real money balances (indicating a higher demand for broad money, the higher the own rate of return), whereas higher long-term rates are negatively related to real money

[14]The short-term interest rate refers to the three-month T-bill rate. Although a preferable measure of the own rate of return would be the actual bank deposit rate, such a series exists only since 1978. Nevertheless, the T-bill rate seems to be a good proxy for the own rate of return, as it is highly correlated with the deposit rate; indeed, the correlation coefficient between the two interest rate series is 0.95 for the period 1978–98.

[15]In fact, Podivinsky (1998) shows that it is preferable to overspecify the number of variables in the model and later add exclusion restrictions, rather than underspecifying the model, as the latter has low power in detecting the true number of cointegrating vectors.

[16]The restricted cointegrated vectors in Figure 2 are given by the fourth specification in the lower part of Table 4, that is, CV_{ppp} = [*p* + 0.88**e* – 1.28**prow*] and CV_{md} = [*p* – *m*3 + 1.22**y* – 0.04**i-long* + 0.02**i-short*].

Table 3. Cointegration Analysis of PPP and Demand for Broad Money

Rank	*Eigenvalue*	*Lambda*	*Critical Value (95%)*	*Trace*	*Critical Value (95%)*
r = 0	0.48	72.34**	45.3	189.6**	124.2
r <= 1	0.32	42.72*	39.4	117.2**	94.2
r <= 2	0.28	36.71*	33.5	74.49*	68.5
r <= 3	0.20	24.24	27.1	37.79	47.2
r <= 4	0.07	8.02	21.0	13.54	29.7
r <= 5	0.05	5.35	14.1	5.53	15.4
r <= 6	0.00	0.18	3.8	0.18	3.8

Restricted Cointegrating Vectors[1]							Joint Test Chi-Square[2]
p	*e*	*q*	*m*3	*y*	*i-long*	*i-short*	
1	**0.95**	**–1.19**	0	0	0	0	0.88
1	0	0	**–1.07**	**1.72**	**–0.02**	**0.02**	
1	**1.06**	–1	0	0	0	0	6.02
1	0	0	**–1.09**	**1.88**	**–0.02**	**0.02**	
1	1	–1	0	0	0	0	7.26
1	0	0	**–1.08**	**1.83**	**–0.02**	**0.02**	
1	**0.88**	**–1.28**	0	0	0	0	6.57
1	0	0	–1	**1.22**	**–0.04**	**0.02**	
1	**0.88**	**–1.34**	0	0	0	0	7.91
1	0	0	–1	1	**–0.04**	**0.02**	
1	**0.89**	**–1.52**	0	0	0	0	12.72*
1	0	0	–1	0.5	**–0.05**	**0.02**	
1	1	–1	0	0	0	0	17.74**
1	0	0	–1	1	**–0.03**	**0.03**	

Notes: Endogenous variables: [*p*, *e*, *q*, *m*3, *y*, *i-long*, *i-short*], *four lags included*
Unrestricted variables: [*Seasonal dummies*, *Time-dummy* 1994:1–1998:2]
Time period: 1971:1–1998:2 (*110 observations*)

[1]The estimations assume three cointegrating vectors, where the third vector is unconstrained. Bold figures are estimated coefficients.

[2]* and ** indicate rejection of the joint likelihood ratio test at the 5 percent and 1 percent significance level, respectively.

balances (indicating a lower demand for money, the higher the rate of return on the alternative asset). Moreover, constraining the coefficient on *m*3 to –1 yields an estimated income elasticity of 1.22. This coefficient is not significantly different from 1 but is significantly different from 0.5. Hence, the long-run income elasticity for broad money seems to be greater than 0.5 but not significantly different from unity.

Recursive estimations of the long-run parameters (the restricted cointegration vectors) further show that the estimated coefficients are quite stable. Figure 3 plots recursive estimates of the coefficients for the nominal effective exchange rate, *e*,

Figure 2. Restricted Cointergration Vectors

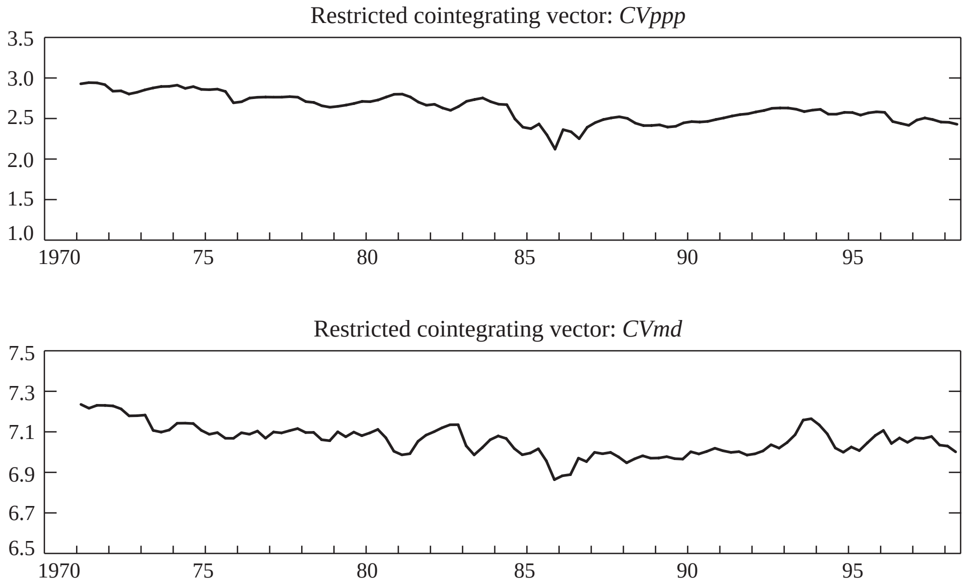


Table 4. Weak Exogeneity Tests

Restricted Cointegrating Vectors (β matrix)

p	*e*	*q*	*m*3	*y*	*i-long*	*i-short*
1	0.88	–1.28	0	0	0	0
1	0	0	–1	1.22	–0.04	0.02
0.99	0.32	–0.90	–0.68	2.41	0.01	0.01

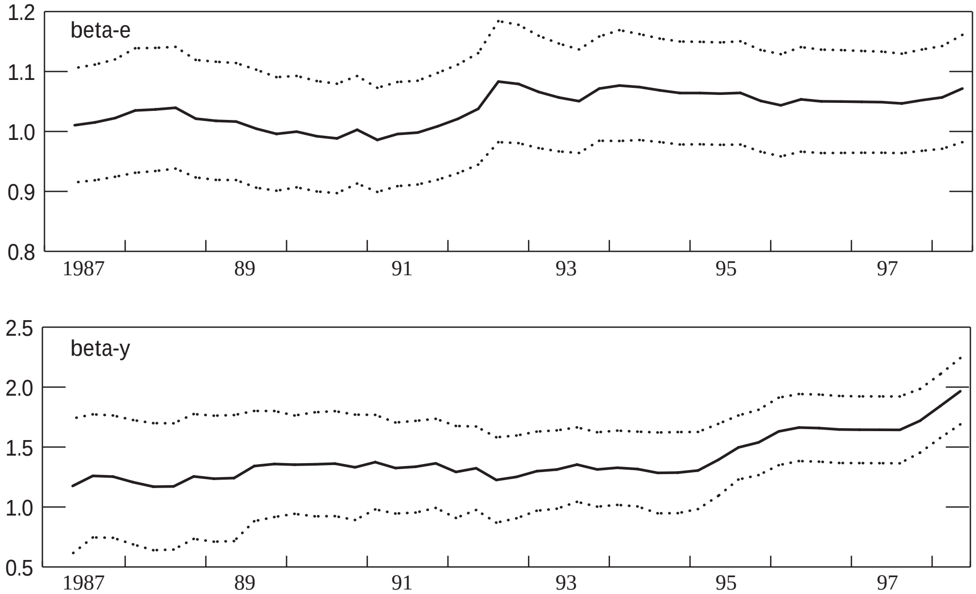
Adjustment Matrix (α matrix)[1]

p	*e*	*q*	*m*3	*y*	*i-long*	*i-short*
–0.05	–0.25	0	–0.09	0.04	2.43	5.88
(0.02)	(0.10)	—	(0.03)	(0.02)	(1.29)	(1.50)
0.01	0.33	0	0.11	–0.02	1.34	–10.08
(0.03)	(0.18)	—	(0.05)	(0.03)	(2.29)	(2.66)
–0.03	–0.23	–0.05	0.15	–0.17	–4.31	6.96
(0.04)	(0.20)	(0.01)	(0.06)	(0.04)	(2.71)	(3.14)

Only restrictions on β:	Chi-sq (3): 6.57
Additional α restrictions:	
q (CV_{md} *and* CV_{ppp}) = 0	Chi-sq (2): 0.00
p (CV_{ppp}) = 0	Chi-sq (1): 14.18**
e (CV_{ppp}) = 0	Chi-sq (1): 4.94*
p (CV_{md}) = 0	Chi-sq (1): 0.24
*m*3 (CV_{md}) = 0	Chi-sq (1): 2.89
y (CV_{md}) = 0	Chi-sq (1): 0.18
i-long (CV_{md}) = 0	Chi-sq (1): 0.29
i-short (CV_{md}) = 0	Chi-sq (1): 5.26*

[1]Standard errors in parentheses. The symbols * and ** indicate rejection of the test at the 5 percent and 1 percent significance level, respectively.

Figure 3. Recursive Cointegration Results
(point estimates and +/–2 standard errors)



and real income, y, including ± 2 standard errors. The recursive initial lag length was chosen to 65, leading to a minimum time span of about 16 years. The coefficient on q was constrained to –1 in the PPP relationship, implying that the coefficient on e is an estimate of the long-run relationship between the nominal effective exchange rate and price differentials. Although there are indications of a small break in the relationship around 1992–93, the estimated coefficients of e are remarkably stable around 1. This further strengthens the result that the PPP hypothesis is a reasonable description of the long-run relationship between developments in the nominal exchange rate and inflation differentials. With regard to the recursively estimated coefficient on y, it is fairly stable around 1.5, although there are indications of an upward trend break during the late 1990s.[17]

Short-Run Dynamics

Thus far, the results show that the variables in the model tend to move together in the long run as predicted by economic theory. To draw policy conclusions, however, the issues of whether the variables should be treated as exogenous or endogenous and how they interact in the short run become important. More generally, it is often useful to examine how the economy adjusts toward the long-run

[17]The recursive estimations occasionally failed to converge due to the sharp reductions in the number of observations. Hence, these estimations were done under the assumption of two (rather than three) cointegrating relationships, and without constraining the coefficient on $m3$ to –1. Nevertheless, the estimated coefficient on $m3$ was always relatively close to –1; the fact that it was not constrained to this value implies that the recursively estimated coefficient on y is a somewhat (upward) biased estimate of the income elasticity.

equilibria—here described by the money demand and PPP relationships—following various types of shocks. These issues are discussed in this section.

The exogeneity issue is addressed by adding exclusion restrictions on the α-matrix in equation (4) in the preferred model (that is, the fourth specification in the lower part of Table 3).[18] The results indicate that foreign prices are clearly exogenous (see Table 4); as expected, foreign prices do not adjust to any disequilibrium in the South African markets. With regard to the PPP relationship, it can be noted that both domestic prices and the nominal exchange rate should be treated as endogenous, as both variables adjust following deviations from the long-run PPP equilibrium. Although the magnitude of these coefficients may seem small, they are actually larger than what is typically found in other countries. In particular, the estimated coefficient of –0.25 for the *e* coefficient in the PPP equation implies that one quarter of any deviation from PPP is adjusted for by a movement in the nominal exchange rate within one quarter, that is, the half-life of such a shock is roughly one year.[19]

In contrast, domestic prices (together with real income and long-term interest rates) could be treated as weakly exogenous in the money demand relationship. The adjustment process—following a deviation from the estimated long-run equilibrium in the money market—seems rather to come through the short-term interest rate and, to some extent, through a change in money holdings.[20] The negative sign on the *i-short* coefficient in the money demand equation implies that a positive shock to real money balances leads to higher short-term interest rates in the following quarter, possibly reflecting a deliberate tightening of monetary policy by the Reserve Bank.[21]

An alternative (and, arguably, more informative) approach to studying the short-run dynamics and comovements among the variables is to examine the impulse response functions from the structural VAR model. This approach allows us to investigate the impact of different types of shocks on both the variables in the model and the estimated equilibrium relationships. In addition, the impulse response functions give us an indication of the lag structure in the economy, which could be useful, for example, from an inflation forecasting perspective. Hence, impulse response functions with an eight-year horizon (32 quarters) were generated after shocking the *p*, *e*, *q*, *m*3, and *y* equations, respectively. Each innovation was obtained

[18]As discussed in the first section, the coefficients in the α-matrix capture the speed of adjustment of a particular variable to a deviation from the long-run equilibria; thus, a zero restriction on any coefficient in this matrix correspond to the null hypothesis that the particular variable does not adjust to restore the long-run equilibrium, and therefore can be treated as weakly exogenous.

[19]As a comparison, MacDonald (1995) finds that the average speed of the nominal exchange rate adjustment following a deviation from PPP is about 2 percent per month for a set of bilateral U.S. dollar exchange rates, implying a half-life of a shock to PPP of about 36 months.

[20]The Chi-square statistic of 2.89 for *m*3 in Table 4 is significant at the 10 percent level.

[21]The main monetary policy instrument in South Africa is the overnight interest rate rather than *i-short* (the T-bill rate). In practice, however, the T-bill rate closely tracks fluctuations in the overnight rate (the correlation coefficient between the two interest rate series is 0.99 between 1970–98), and could therefore be regarded as controlled by the Reserve Bank.

[22]The practical significance of the ordering is that a shock to a variable is allowed to have contemporaneous effects only on the variable itself and the succeeding variables in the ordering. Thus, the assumed ordering implies that a shock to, for example, real output may have a contemporaneous effect on the nominal variables *m*3, *p*, *e*, and *i*, while a shock to any of these nominal variables can only affect real output with (at least) a one quarter lag. In the current study, the reported results are quite robust to alternative orderings of the variables.

by a standard Choleski decomposition, where the ordering of the variables, in general, matters. Somewhat arbitrarily the chosen ordering was *q*, *y*, *m*3, *p*, *e*, *i-short*, and *i-long*.[22]

The results are illustrated in Figures 4 and 5. To start with, a deviation from the long-run PPP relation can occur due to a shock to the nominal exchange rate, domestic prices, or foreign prices. The impacts of shocks to these equations are shown in Figure 4; the left-hand panels show the adjustments over time of some selected individual variables, while the right-hand panels show the developments of the deviations from the estimated long-run PPP equilibrium.

A positive shock to domestic prices will lead temporarily to an appreciation of the real exchange rate. The nominal exchange rate, however, will start to depreciate sharply already after three to four quarters, peaking after about eight quarters. In fact, the response of the exchange rate will be sufficiently sharp to cause an overshooting effect leading to a temporary real depreciation before equilibrium is restored, as illustrated in the dynamic effects on the cointegrating vector.

Likewise, a shock to the nominal exchange rate, say a depreciation (a negative shock), will almost immediately result in higher inflation. The half-life time of such a shock seems to be about five quarters, a result that is in line with the earlier result from the α-matrix. The effect on real output from a shock to the *e* equation is plotted in the middle (left) panel. It is interesting to note that, although a shock to the nominal exchange rate leads to a quite persistent effect on the real exchange rate (the real exchange rate remains, say, depreciated for several years before full adjustment takes place), the impact on real output is virtually zero.

Finally, a shock to foreign prices will cause domestic prices to respond (as earlier noted). Nevertheless, the real exchange rate will temporarily deviate from the long-run equilibrium, as the nominal exchange rate also adjusts. The response of domestic prices peaks after about eight quarters, before the variables settle on a new path to restore equilibrium.

Turning to the short-run comovements of the variables in the money demand function, the left-hand panels in Figure 5 show the dynamic effect of a one standard error shock to the *m*3 equation. Again, it should be noted that a shock to money can originate from different sources and should not necessarily be interpreted as a monetary policy shock. Nevertheless, it is interesting that, in contrast to a shock to the nominal exchange rate, a positive shock to money leads to an initial but temporary output gain that peaks after about one year. The excess money balances also seem to trigger a tightening of monetary policy, however, as the short-term interest rate picks up sharply after a couple of quarters. These effects also imply that the cointegrating relationship is driven back toward its equilibrium as higher real balances are offset by higher output and higher short-term interest rates. Domestic prices start to pick up after about five to six quarters, implying that real money balances adjust back toward their initial level at the same time as the output effect vanishes and equilibrium is restored. One can also notice that long-term interest rates pick up after a couple of quarters, indicating that inflation expectations (correctly) rise.

Finally, a positive shock to real output leads quickly to higher demand for real balances and holding of broad money rises. The expected impact on domestic

Figure 4. Impulse Responses of Shocks to the PPP Relation

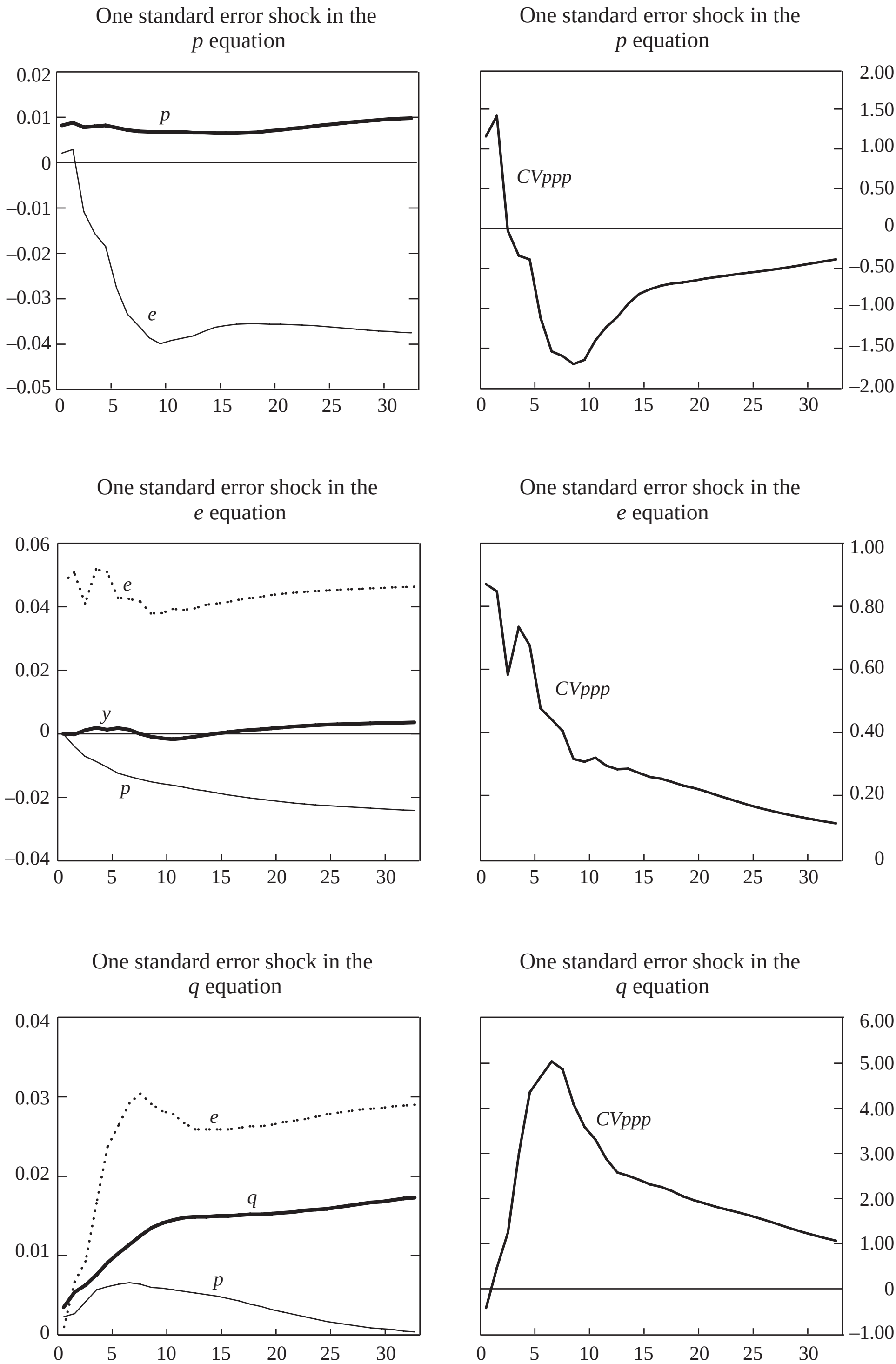
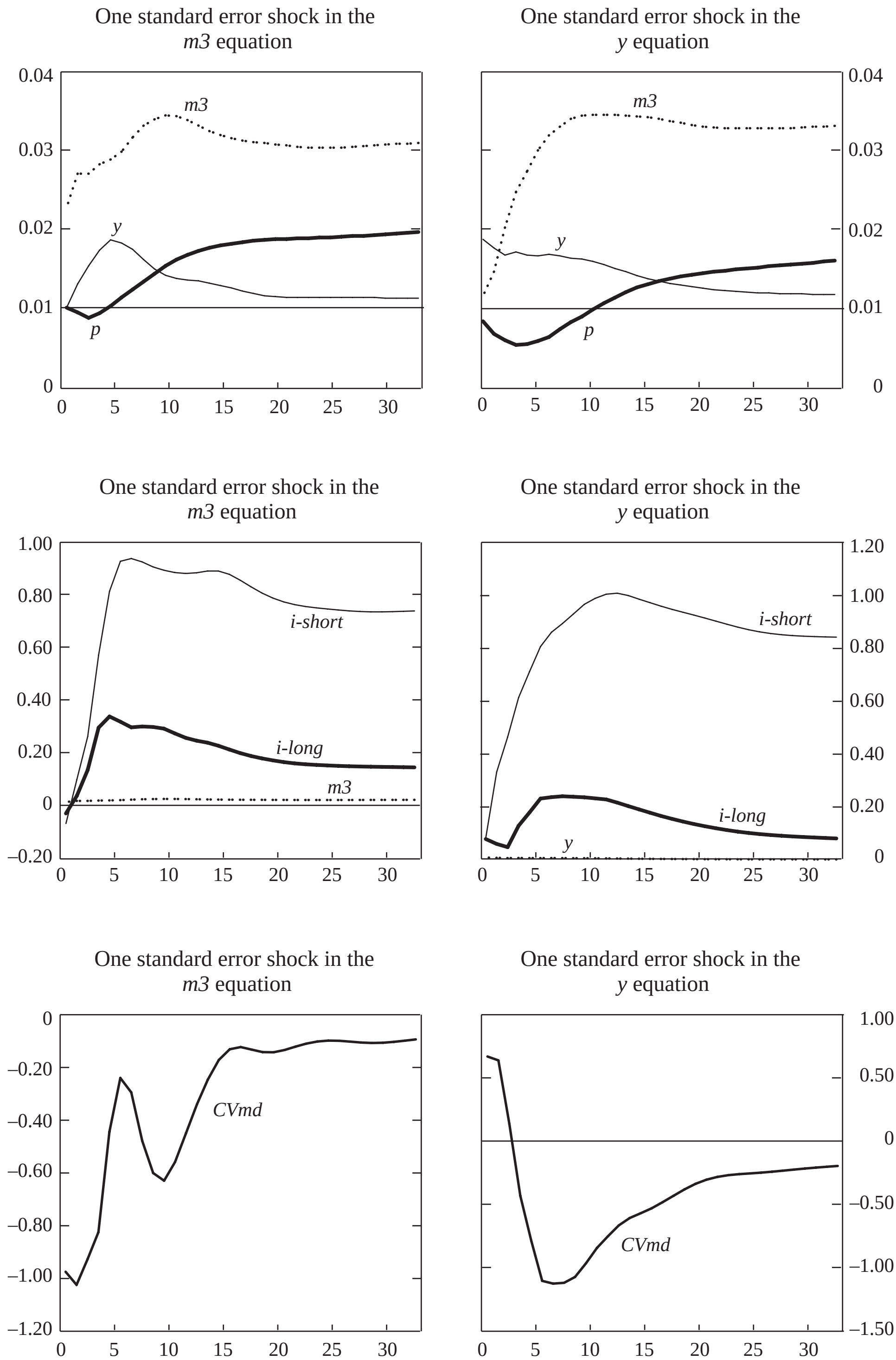


Figure 5. Impulse Responses of Shocks to the Money Demand Relation



prices is in principle ambiguous, as it depends on whether the output shock is driven by a shift in aggregate demand or aggregate supply. The empirical results indicate, however, that a positive shock to output results in inflationary pressures after about four to five quarters; although there will initially be some downward pressure on domestic prices, the end effect is a higher price level. Again, the magnitude of these inflationary pressures seems to be mitigated by a tightening of monetary policy, as short-term interest rates rise.

The above results are similar to what is found in a number of other countries. For example, Sims (1992) uses a similar VAR setup to study the effects of monetary policy in five OECD countries. He shows that a shock to the money equation results in a temporary real output response in France, the United Kingdom, and the United States, while inflation adjusts with a lag. He also shows that a positive shock to real output results in upward pressure on domestic prices in several countries. Likewise, Eichenbaum (1992) shows that a positive shock to money (M1) results in a small and temporary increase in output in the United States, but also in a sharp and substantial adjustment in the short-term interest rate (the federal funds rate), while domestic prices pick up gradually and peak after about four to five quarters.

III. Discussion and Conclusions

The results in this paper indicate: (i) a stable money demand type of relationship exists among domestic prices, broad money, real income, and nominal interest rates, with plausible estimates of the long-run coefficients, as well as a long-run relationship among domestic prices, foreign prices, and the nominal effective exchange rate; and (ii) in the short run, shocks to the exchange rate affect domestic prices but have virtually no impact on real output, while shocks to broad money have a temporary impact on real output before inflation picks up. Both types of shocks seem to trigger a monetary policy response, as the short-term interest rate adjusts quickly and substantially.

An interesting aspect of the results is that even though the South African economy has undergone a number of important structural changes during the studied period—including long periods of trade sanctions, the presence of the financial rand system and widespread exchange controls on residents, different monetary policy regimes (see Box 1), and considerable swings in the terms of trade—the long-run relationships among the examined macroeconomic and financial aggregates are fairly stable and consistent with economic theory. In this context, it is perhaps not surprising that it is the broadest measure of money that seems to work better in the long run, as the more narrow money aggregates possibly would exhibit more frequent structural breaks.

The result with regard to the PPP relationship is, perhaps, somewhat surprising. Although considerable evidence for other countries suggests that the real exchange rate is mean-reverting in the long run, especially for small open economies with floating exchange rates (see MacDonald and Marsh, 1997; and Rogoff, 1996), homogeneity and symmetry restrictions are often rejected, and deviations from PPP tend to dampen out only at a relatively slow rate. Also, as noted earlier, Aron, Elbadawi, and Kahn (1997) find that the real exchange rate in South Africa is non-stationary but cointegrated with a set of "fundamentals."

In the current study, the results clearly show that domestic prices, foreign prices, and the nominal exchange rate form a cointegrated vector in South Africa. Moreover, the estimated coefficients of this vector are relatively close to unity (in absolute terms), although a strict test of PPP is rejected in some specifications. When comparing these results to the ones in Aron, Elbadawi, and Kahn (1997), it is important to note that the current study allows for a structural break in 1994.[23] This break is intended to capture the opening up of the South African economy in recent years, which have included a significant degree of trade liberalization (see Jonsson and Subramanian, 2000) as well as financial deepening. Indeed, Aron, Elbadawi, and Kahn (1997) find that various measures of trade liberalization is an important determinant of the real exchange rate and note, for example, that the reduction of tariffs since 1994 is consistent with a more depreciated real exchange rate. These findings would thus be in line with the results in the current study.

Aron, Elbadawi, and Kahn (1997) also argue that the South African Reserve Bank appears to have actively stabilized the real exchange rate during certain periods, in which case the PPP finding in the current study rather should be interpreted as a policy reaction function. It should also be emphasized that the findings in this paper do not preclude deviations in the real exchange rate from its equilibrium level from time to time with potentially damaging effects on South Africa's competitiveness.

With regard to the money demand relationship, it is interesting to note that notwithstanding the structural changes mentioned above, the magnitude and pattern of the estimated relationship is similar to what is found in many other industrial countries. Both the estimation of a long-run income elasticity close to 1 and the estimated short-run comovements among the variables are in line with the empirical results found for many other countries, see Fase (1993) and Sims (1992), respectively. For example, the results that shocks to either money or output affect domestic prices with a lag of four to six quarters is similar to what is found in several other inflation-targeting countries. Taken together, these results suggest that it would be possible to develop a satisfactory forecasting model for inflation in South Africa, which is similar to the ones used in other countries that have adopted an inflation-targeting framework for monetary policy.

[23]It should also be noted that the definitions of domestic prices are different (the current study uses "underlying CPI," whereas Aron, Elbadawi, and Kahn (1997) use wholesale price index) and that the examined time period is somewhat different.

APPENDIX

Unless otherwise indicated, the data series are from the South African Reserve Bank (SARB), *Quarterly Bulletin*.

p: Underlying consumer price index. This index was provided by the SARB for 1975–98, and equals headline consumer price index excluding "food and non-alcoholic beverages," "home owner's cost" and "value-added tax." For 1970–75, the series was defined as the headline consumer price index net of food prices.

nc: Notes and coins outside the banking system.

***m*3**: *nc* plus checking deposits, and short-, medium-, and long-term deposits.

i-short: Interest rate on three-month T-bills. Source: *International Financial Statistics* (*IFS*), IMF.

i-long: Interest rate on ten-year government bonds. Source: *IFS*, IMF.

e: Nominal effective exchange rate including (weights in brackets) U.S. dollar (51.7), pound sterling (20.2), deutsche mark (17.2), and Japanese yen (10.9).

q: Effective consumer price index in foreign countries, including the same four countries and weights as when calculating *e*. Source: *IFS*, IMF.

y: Gross domestic product, 1990 prices, seasonally adjusted.

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