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

Prepared by Gunnar Jonsson[1]

Authorized for distribution by Michael Nowak

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Abstract

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, since the short-term interest rate adjusts quickly.

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I. Introduction

Inflation has fallen considerably in South Africa 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. However, developments in broad money (M3) have not followed the same pattern. Although the annual growth rate in M3 fell from 14 percent in 1991 to 5 percent in 1993, it increased to 17 percent in 1998. The contrasting developments in inflation and money growth during the 1990s have led analysts to question whether there exists a stable relationship between money growth and inflation, i.e., whether money demand is stable, and whether it is appropriate for the Reserve Bank to use M3 as an intermediate target for its monetary policy decisions.[2] [3]

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 per year between 1990-95, but by 22 percent and 19 percent in 1996 and 1998, respectively. The latter events were followed by a pick-up in inflation, despite a tightening of monetary policy. In general, it is likely that movements in foreign prices and the nominal exchange rate have contributed to inflation developments in South Africa also during the 1970s and 1980s, as South Africa remained a fairly open economy during this period, notwithstanding long periods of international trade and financial sanctions.[4]

The purpose of this empirical study is twofold: First, the study examines the long-run stability of two economic relationships involving domestic prices—a money demand type of

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[2] Strictly speaking, an unstable relationship between money growth and inflation does not necessarily imply that money demand is unstable, as the latter would be expected to vary with fluctuations in other variables, such as real income and nominal interest rates.

[3] The Reserve Bank has announced annual guidelines for growth in broad money since 1986 (6-10 percent in 1997). However, actual growth in M3 has substantially exceeded the guideline range during the last few years, and the authorities have on several occasions announced that the Reserve Bank in practice is guided by developments in a number of different indicators, including various price indices, the shape of the yield curve, the nominal exchange rate, and the output gap (see South African Reserve Bank (1998)).

[4] 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, e.g., Jonsson and Subramanian (1999) and Lipton (1998) for discussions. The authorities also operated a dual exchange rate system and had extensive capital controls in place during long periods up to 1995, with financial transactions typically being traded at a discounted exchange rate (the financial rand mechanism); see Garner (1994) for a description of this system. In the current paper, the market determined exchange rate for the commercial rand is used throughout the study, see Appendix for details.

relationship and purchasing power parity (PPP). Second, 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.

From a methodological perspective, an advantage with the current study is 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; however, as both relationships involve domestic prices, it seems preferable to model the interaction among the variables within a multivariate cointegration context.[5]

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.

The rest of this paper is organized as follows. Section II briefly reviews some earlier studies related to inflation, money demand, and PPP in South Africa, before it lays out the theoretical and methodological approach in the current study, and discusses some data issues. Section III presents the empirical results, and Section IV summarizes the findings and concludes.

II. Background, Methodology, and Data

A. 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.[6] This is also the case for South Africa, where previous empirical studies on the relationship among prices, money, and the exchange rate have either focused exclusively on the demand for money or the degree of pass-through from exchange rate movements to inflation. The most recent studies on the demand for

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[5] 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.

[6] 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. McDonald (1995), Rogoff (1996), Nessén (1996) and Habermeier and Mesquita (1999) are examples of studies that survey the PPP literature and provide some new results using cointegration methods.

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.[7] 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.

The issue of PPP and whether there is a full pass-through from exchange rate movements to domestic prices in South Africa have been discussed by, e.g., Tsikata (1998) and Subramanian (1998). Although the results are sensitive to the choice of price aggregates and sample period, indications are that the effective nominal depreciation of the rand during the 1990s is almost fully reflected in higher prices of imported goods.

B. Theoretical Background

The theoretical underpinning for the study of money demand and PPP is standard. Hence, assume that the domestic price level in the economy, p, is a linear combination of the price level for tradeable goods, p_T, and non-tradeable goods, p_{NT}, respectively, i.e.,

$$p = \theta p_T + (1-\theta) p_{NT}, \qquad (1)$$

where $0 < \theta < 1$. The PPP hypothesis implies that the price of the tradeable good is determined in the world market and equal to (in logs)

$$p_T = q - e \qquad (2)$$

where e is the nominal (effective) exchange rate (defined as foreign currency per domestic currency), and q represents (effective) foreign prices. The price of non-tradeables is then determined by adding an assumption of equilibrium in the money market; real money supply $(m^s - p)$ equals real money demand, with the latter assumed to being a positive function of real income, y, and a negative function of the nominal interest rate, i,[8] i.e.,

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[7] Earlier money demand studies for South Africa include Courakis (1984), Whittaker (1985), Tavlas (1989), and Hurn (1991). However, as pointed out by Hurn and Muscatelli (1992), these studies do not consistently estimate the long-run elasticities of the variables in the money demand function.

[8] 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 broader money typically is interest bearing. See Section III.A for further discussion.

$$m^s - p = m^d(y, i) \qquad (3)$$

Thus, we end up with a system of 6 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 arguably have 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.

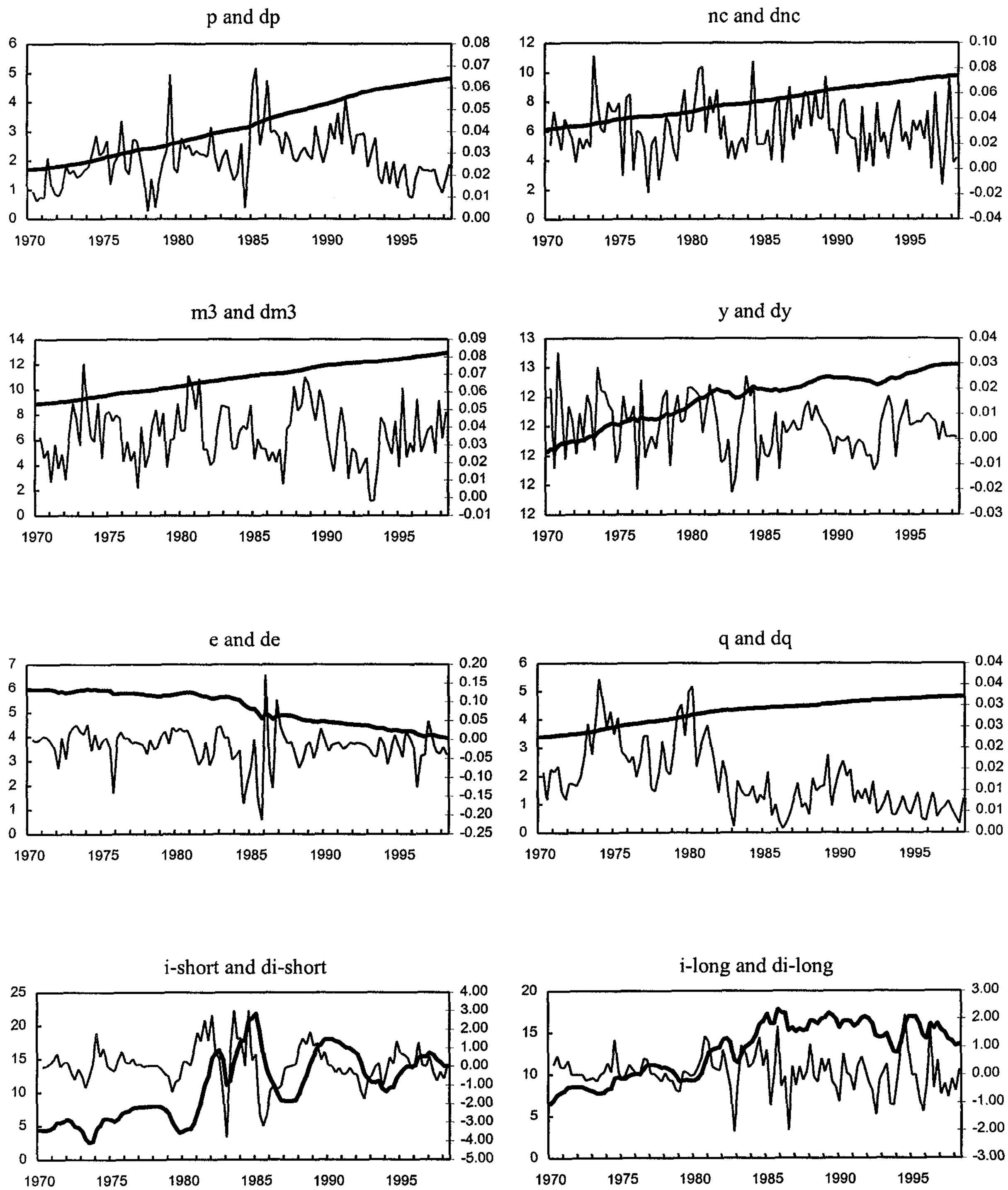
C. Data Issues and Econometric Methodology

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 *m3* (broad money), consisting of *nc* plus cheque 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 non-stationary in levels but stationary in first-differences.[9] The non-stationarity of the data together with the notion that

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

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 LHS; first differences on RHS.

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
m3	-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
dm3	-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; * indicate rejection of the null hypothesis of non-stationarity at the 5-percent significance level.
2/ The lag length was chosen by using the Schwarz Bayesian Criterion assuming a maximum of 8 lags.
3/ Including a time trend in the first difference regressions did not alter the results qualitatively, except for *q* for which non-stationarity was rejected if a trend was included.

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 there existed any long-run relationship among the variables.[10] 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,

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

$$\Delta x_t = \mu + \sum_{i=1}^{k} \Gamma_i \Delta x_{t-i} + \pi x_{t-1} + \varepsilon_t \qquad (4)$$

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 (i.e., when rank(π) $= r < n$, 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 equlibria (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

$$\pi x_{t-1} = \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \\ \alpha_{41} & \alpha_{42} \\ \alpha_{51} & \alpha_{52} \\ \alpha_{61} & \alpha_{62} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{12} & \beta_{13} & \beta_{14} & \beta_{15} & \beta_{16} \\ \beta_{21} & \beta_{22} & \beta_{23} & \beta_{24} & \beta_{25} & \beta_{26} \end{bmatrix} \begin{bmatrix} p \\ q \\ e \\ m \\ y \\ i \end{bmatrix}_{t-1} . \qquad (5)$$

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 $[\beta_{14} \;\; \beta_{15} \;\; \beta_{16}]$ and $[\beta_{22} \;\; \beta_{23}]$, respectively.

The short-run comovements among the variables are then examined by generating orthogonalized impulse response functions (see Sims (1980)). However, as opposed to the traditional VAR literature, the computation of the impulse response functions is based on the VECM representation (expression (4)), 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 are on shocks to the money, exchange rate, price, and output equations, respectively, in (5), 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, e.g., Becker (1999) for a discussion of these identification issues.

III. Results

A. 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.[11] 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 4 lags in the VAR when quarterly data are used, at this stage, the results are reported with 2, 3, and 4 lags included in the VAR, respectively. Formal tests for the appropriate lag length are reported later.[12]

The economic model suggests that 2 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 column 2 in Table 2). However, the results varied between 0 and 4 vectors 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 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 2 of the 12 specifications (column 3, Table 2),[13] and the hypothesis that one cointegrating vector includes only the variables p, m, y and i was rejected in 4 of the 12 specifications (column 4, Table 2). However, the joint test of the two hypotheses was rejected 8 of the 12 specifications.

Turning to the parameters of the cointegrating vectors,[14] 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.

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[11] The number of cointegrating vectors were 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.

[12] 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.

[13] By not constraining the coefficients on q and e to equal 1 (in absolute values), the test allows for various fixed costs, such as transportation and menu costs, to vary over time and across countries. This test is sometimes referred to as a simple test of PPP.

[14] 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 Section II.B.

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, *m3*.						
					p	*e*	*q*	*m3*	*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 2 cointegrating vectors. Bold figures are estimated coefficients.

Indeed, it is interesting to notice 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 one 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, *m3*, is included in the model. The estimated coefficient on *m3* 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 become 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 *m3* 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. Hence, the model which included *m3* was re-estimated 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 money, while *i-long* would measure the rate of return of alternative assets. A test for system reduction of the number of lags in the VAR indicated that 4 lags should be included (see Table 3). The cointegration tests from this model are reported in Table 4. 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 relationship discussed above, it is conceivable that the two interest rates together form another long-run relationship—possibly together with domestic prices and/or the exchange rate; the theory of the term structure of interest rates suggests that a linear combination of long-term and short-term interest rates contain information about the future monetary policy stance and could help in extracting expectations about future inflation rates and/or nominal exchange rate movements (see, e.g., 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]

[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.

Table 3. Tests for System Reduction of Lags in the VAR

Lag-reduction	*F*-value
5 lags to 4 lags:	*1.16*
6 lags to 4 lags:	*1.44***
7 lags to 4 lags:	*1.23*
8 lags to 4 lags:	*1.09*
4 lags to 3 lags:	*1.43**
5 lags to 3 lags:	*1.30**
6 lags to 3 lags:	*1.48***
7 lags to 3 lags:	*1.30**

Endogenous variables: *[p, e, q, m3, y, i-long, i-short]*

Unrestricted variables: *[Seasonal dummies, Time-dummy 1994:1 - 1998:2]*

Time period: *1972:1 - 1998:2 (106 observations)*

Note: * and ** indicate rejection of the null-hypothesis at the 5-percent and 1-percent significance level, respectively.

The results with regard to the money demand type of relationship and PPP 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 type of relationship and PPP could not be rejected (see Table 4), 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.[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 inflation 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-term and short-term interest rates entered the vector with different and predicted signs,

[16] The restricted cointegrated vectors in Figure 2 are given by the fourth specification in the lower part of Table 4, i.e., *CV-ppp = [p + 0.88*e -1.28*prow]* and *CV-md = [p - m3 + 1.22*y - 0.04*i-long + 0.02*i-short]*

Table 4. Cointegration Analysis of PPP and Demand for Broad Money

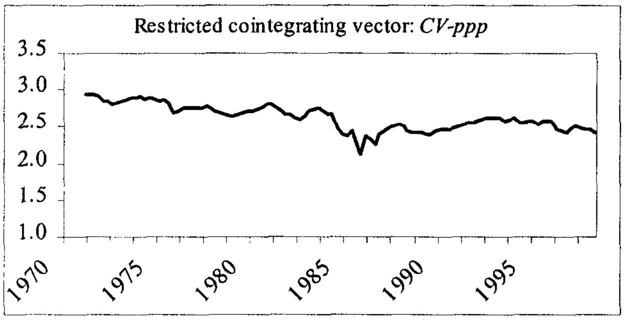
Rank	Eigenvalue	Lambda	Critical value (95%)	Trace	Critical value (95%)
r = 0	0.48	72.34**	45.3	189.6**	124.2
r <= 1	0.32	42.72*	39.4	117.2**	94.2
r <= 2	0.28	36.71*	33.5	74.49*	68.5
r <= 3	0.20	24.24	27.1	37.79	47.2
r <= 4	0.07	8.02	21.0	13.54	29.7
r <= 5	0.05	5.35	14.1	5.53	15.4
r <= 6	0.00	0.18	3.8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

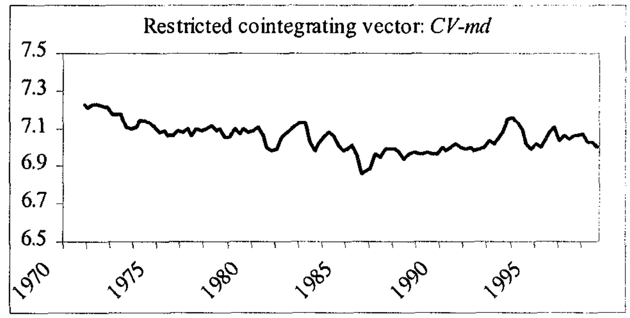
Restricted cointegrating vectors 1/							Joint test 2/ Chi-square
p	e	q	m3	y	i-long	i-short	
1	**0.95**	**-1.19**	0	0	0	0	0.88
1	0	0	**-1.07**	**1.72**	**-0.02**	**0.02**	
1	**1.06**	-1	0	0	0	0	6.02
1	0	0	**-1.09**	**1.88**	**-0.02**	**0.02**	
1	1	-1	0	0	0	0	7.26
1	0	0	**-1.08**	**1.83**	**-0.02**	**0.02**	
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**	

Endogenous variables: *[p, e, q, m3, y, i-long, i-short], 4 lags included*
Unrestricted variables: *[Seasonal dummies, Time-dummy 1994:1 - 1998:2]*
Time period: *1971:1 - 1998:2 (110 observations)*

1/ The estimations assume 3 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.

Figure 2. Restricted Cointegration Vectors



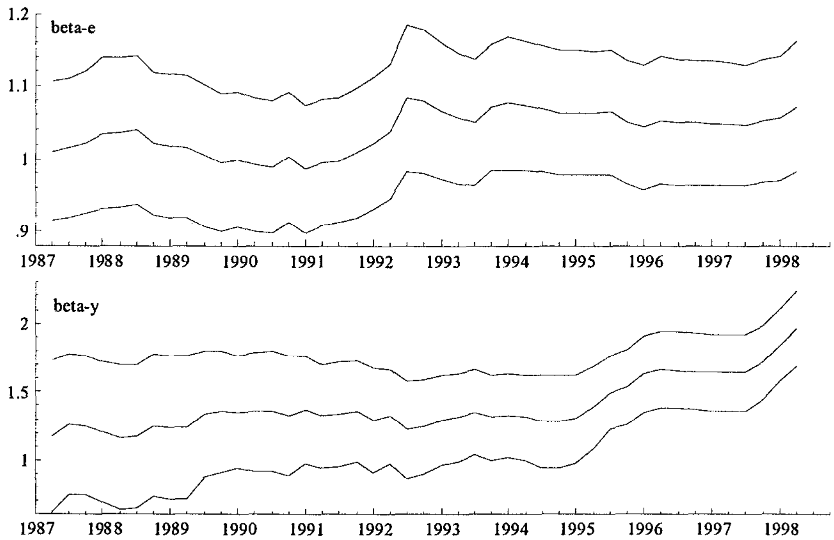


i.e., higher short-term rates are positively related to real money balances (indicating a higher demand for broad money, the higher the own rate of return), whereas higher long-term rates are negatively related to real money balances (indicating a lower demand for money, the higher the rate of return on the alternative asset). Moreover, constraining the coefficient on *m3* to -1 yields an estimated income elasticity of 1.22. This coefficient is not significantly different from 1 but 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*, 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 inflation differentials. This coefficients is remarkably stable around 1, further strengthening the result that the strong PPP hypothesis is a reasonable description of how the nominal exchange rate is related to inflation differentials in the long-run in South Africa. 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]

[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. Despite this, 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.

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



B. 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. However, to draw policy conclusions, 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 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 (5) in the preferred model (i.e., the fourth specification in the lower part of Table 4).[18] The results indicate that foreign prices are clearly exogenous (see Table 5); as expected, foreign prices do not adjust to any disequlibria 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 PPP

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[18] As discussed in Section II.B, 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.

Table 5. Weak Exogeneity Tests

Restricted cointegrating vectors (β matrix)

p	*e*	*q*	*m3*	*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*	*m3*	*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
m3 (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 parenthesis.
2/ * and ** indicate rejection of the test at the 5-percent and 1-percent significance level, respectively.

equilibrium. 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—seem rather to come through the short-term interest rate and, to some extent, through a change in money holdings.[19]

[19] The Chi-square statistic of 2.89 for *m3* in Table 5 is significant at the 10-percent level.

These results are confirmed by examining the error-correction model for the complete system of variables (see Table 6).[20] The dependent variable in each regression is the quarterly percentage change in a particular variable,[21] and the dependent variables are the lagged residuals from the two restricted cointegration vectors—the error-correction terms—denoted *EC-md* and *EC-ppp*, respectively, the time dummy, and lagged observations of all variables in the system. Thus, the estimated coefficients on the error-correction terms pick up the magnitude of the subsequent (quarterly) adjustment of a variable to any disequilibrium in the money market or deviation from PPP.

The results indicate that inflation adjusts following a temporary deviation from the long-run PPP equilibrium; in the event the real effective exchange rate depreciates (above its long-run equilibrium level), underlying inflation will increase in the subsequent quarter to partly offset the real depreciation. However, the adjustment also takes place through a change in the nominal exchange rate, as indicated by the significant coefficient of the *EC-ppp* variable when *de* is the dependent variable. Although the magnitude of these adjustment coefficients may seem small, they are actually larger than what is typically found in other countries. In particular, the estimated coefficient of –0.22 for the *EC-ppp* coefficient in the *de* equation implies that one-fifth of any deviation from PPP is adjusted for by a movement in the nominal exchange rate within one quarter, i.e., the half-life of such a shock is roughly one year.[22]

A disequilibrium in the money market does not immediately trigger an inflation adjustment; rather short-term interest rates adjust strongly in the subsequent quarter to partially offset the disequilibrium. The negative sign on the *EC-md* variable in the *di-short* 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. It can also be noted that the estimated coefficient for the time dummy has the expected negative sign and is significant in the inflation equation, indicating the underlying annual inflation rate after 1993 would be about 4 percentage points less than during 1972-

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[20] The absence of weak exogeneity for a subset of variables implies that it is not appropriate to model a particular variable, such as inflation, in a one-equation error-correcting (parsimonious) model.

[21] Except for the interest rates, where the dependent variable is the change in the interest rate.

[22] As a comparison, McDonald (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.

Table 6. The Error Correction Model, 1971:2 - 1998:2 1/ 2/

	Dependent Variable						
Independent variables	*dp*	*de*	*dq*	*dm3*	*dy*	*di-long*	*di-short*
EC-md (-1)	0.00 [-0.13]	0.42 [2.06]	0.02 [1.31]	0.07 [1.28]	0.04 [1.05]	2.73 [1.04]	12.60 [-4.13]
EC-ppp (-1)	-0.04 [-2.62]	-0.22 [-2.56]	0.00 [-0.20]	-0.08 [-3.20]	0.03 [1.85]	2.09 [1.84]	5.28 [4.01]
dum9498	-0.01 [-2.06]	-0.01 [-0.26]	0.00 [-0.56]	-0.01 [-2.57]	0.01 [2.78]	0.41 [1.59]	0.03 [0.10]

1/ The reported results are for the complete system, with the fourth lag as the only dropped variables. Thus, the system also includes 3 lags of each variable, lagged residuals from the third (unconstrained) cointegrating vector, as well as seasonal dummy variables.
2/ t-statistics in brackets.

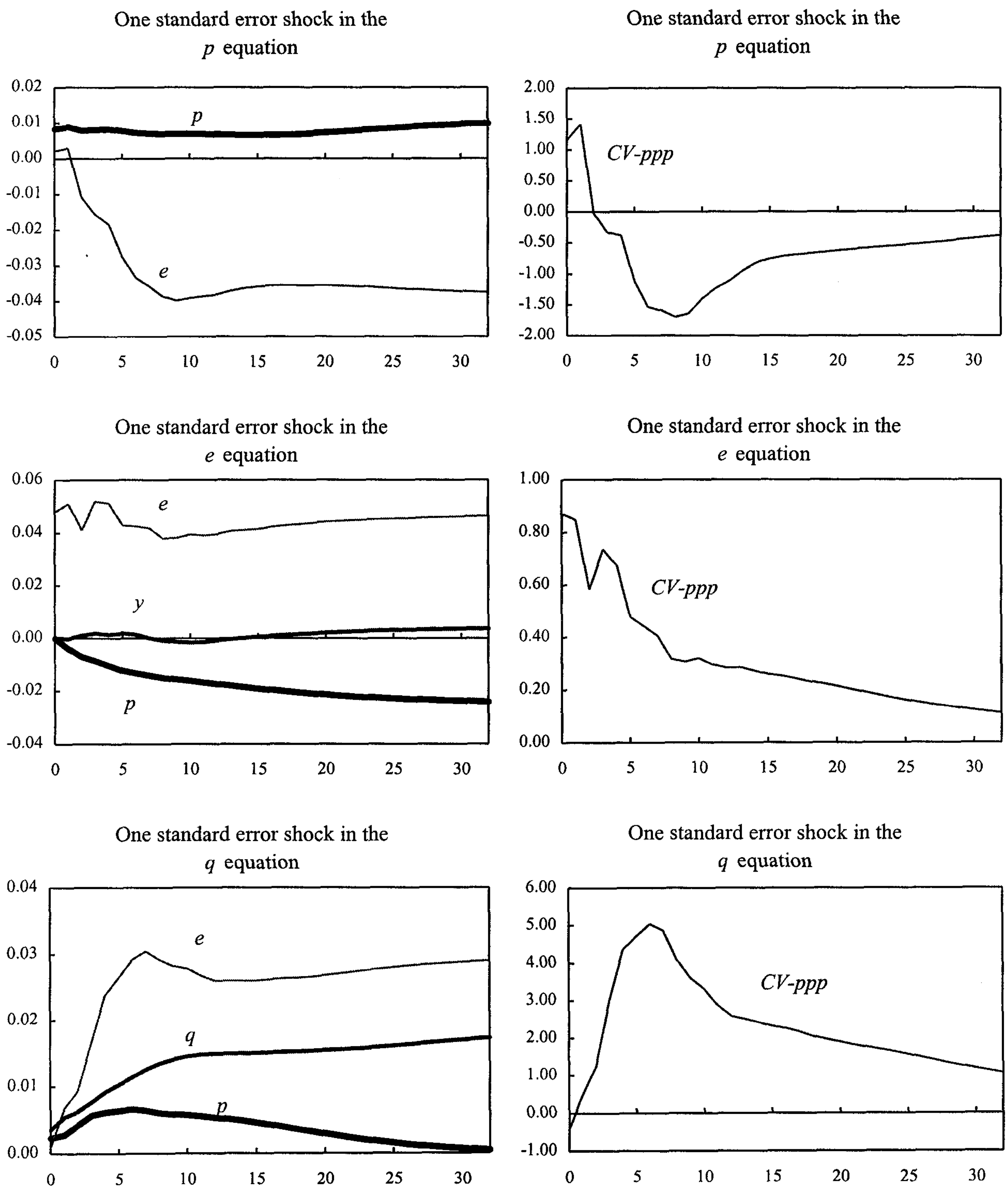
Specification tests
Vector autocorrelation, AR 1-5: F(245,280) = 1.03
Vector normality, Chi-sq (14) = 51.6**
Vector heteroscedasticity, F(1344,309) = 0.35

Tests for system reduction of lags
4 lags to 3 lags: F(49,344) = 1.22
3 lags to 2 lags: F(49,380) = 1.67**
4 lags to 2 lags: F(98,432) = 1.45**

1993, everything else equal, thereby suggesting an important process of financial deepening in recent years.

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 a 8-year horizon (32 quarters) were generated after shocking the *p, e, q, m3,* and *y* equations, respectively. Each innovation was obtained by a standard Choleski decomposition, where the ordering of the variables in

Figure 4. Impulse Responses of Shocks to the PPP Relation



general matters. Somewhat arbitrarily the chosen ordering was *q, y, m3, p, e, i-short, i-long.*[23] 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. However, the nominal exchange rate will start to depreciate sharply already after 3-4 quarters peaking after about 8 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.

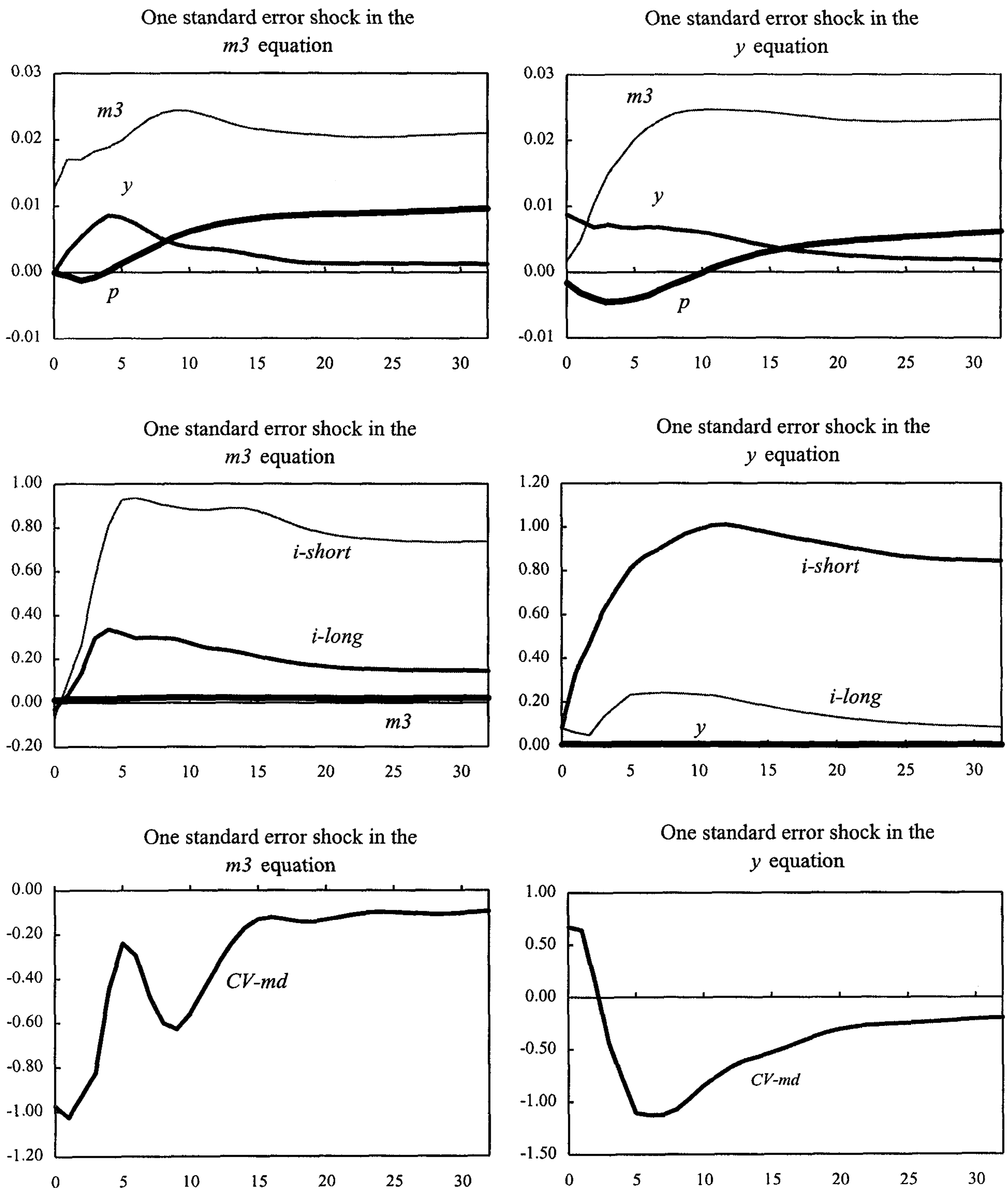
Likewise, a shock to the nominal exchange rate, say a depreciation (a negative shock), will almost immediately result in higher inflation. The half-life time of such a shock seems to be about 6 quarters, a result that is in line with earlier results from the α-matrix and the ECM. The effect on real output from a shock to the *e* equation is plotted in the middle (left) panel. It is interesting to notice 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 noticed). Nevertheless, the real exchange rate will temporarily deviate from the long-run equilibrium, as the nominal exchange also adjusts. The response of domestic prices peak after about 8 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 shows the dynamic effect of a one standard error shock to the *m3* 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 to notice 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 1 year. However, the excess money balances also seem to trigger a tightening of monetary policy, 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 is offset by higher output and higher short-term interest rates. Domestic prices start

[23] 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 *m3, 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.

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



to pick up after about 5-6 quarters implying that real money balances adjust back toward its 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) rises.

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 prices is in principle ambiguous, as it depends on whether the output shock is driven by a shift in aggregate demand or aggregate supply. However, the empirical results indicate that a positive shock to output results in inflationary pressures after about 4-5 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, United Kingdom and the U.S., 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 for the U.S. a positive shock to money (M1) results in a small and temporary increase in output, but also in a sharp and substantial adjustment in the short-term interest rate (the Federal Fund rate), while domestic prices pick up gradually and peaks after about 4-5 quarters.

IV. Discussion

The results in this paper indicate that (i) there exists a stable money demand type of relationship between 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 between 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[24], and considerable swings in the terms of trade—the long-run relationships between the examined macroeconomic and financial aggregates are fairly stable and consistent with economic theory. In this context, it

[24] 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, rather than changes in cash and liquid asset requirements combined with credit ceilings and interest rate controls. This probably enhanced the responsiveness of monetary aggregates to macroeconomic developments.

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.

It is also interesting to notice that despite the structural changes mentioned above, the magnitude and pattern of the estimated relationships are similar to what is typically found in many other industrial countries. Both the estimations of the long-run relationships—in particular, an income elasticity close to 1 in the money demand function, and a long-run coefficient close to 1 for the nominal effective exchange rate in the PPP estimations, see Fase (1993) and Habermeier and Mesquita (1999), respectively—and the estimated short-run comovements among the variables (see Sims (1992)) are in line with the empirical results found for many other countries.

The South African authorities are contemplating the adoption of a formal inflation targeting framework for monetary policy (see South African Reserve Bank (1999)). Given the forward looking nature of such a regime, inflation forecasting is one of the key elements in this framework. The results in the current paper suggest that it would be possible to develop a satisfactory forecasting model for inflation which is similar to the ones used in other countries that have implemented this regime, e.g., United Kingdom. In particular, the results that shocks to either money or output affect domestic prices with a lag of 4-6 quarters is similar to what is found in several other inflation-targeting countries.

Regardless of monetary policy framework in South Africa, two general policy conclusions emerge from the study: (i) The result regarding the stability of money demand means that growth rates in broad money cannot be ignored when monitoring and forecasting inflation developments. (ii) The result regarding the long-run PPP hypothesis together with the lack of a significant short-run impact on real output from nominal exchange rate shocks imply that there is a natural limit to the effectiveness of exchange rate policy; any attempts to enhance competitiveness by depreciating the nominal exchange rate will be dissipated through higher domestic inflation.

Data 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.

m3: *nc* plus checking deposits, and short- medium- and long-term deposits.

i-short: Interest rate on 3-months t-bill. Source: *International Financial Statistics* (IFS), IMF.

i-long: Interest rate on 10-year government bonds. Source: IFS, IMF.

e: Nominal effective exchange rate including (weights in brackets) U.S. dollar (51.7), British pound (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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