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## **Bilateral Trade Elasticities : Sweden Versus Her Major Trading Partners\***

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### **ABSTRACT**

*This study explores the long-run bilateral trade elasticities between Sweden and its six major trading partners for the period 1960-1999. Tests for unit roots and cointegration in a panel perspective are conducted. The estimated cross sectional trade elasticities show that trade is highly sensitive to changes in income but less sensitive to real exchange rate fluctuations. The bilateral trade elasticities disclose that the Marshall-Lerner condition is not satisfied (except for Germany) and real depreciation of the Swedish currency has less favorable impact on the trade balance. The policy implications of our findings are also discussed.*

**JEL classification codes:** F41; F43; C12

**Key words:** *bilateral income elasticity; bilateral price elasticity; the Marshall-Lerner condition; panel unit root; panel cointegration,*

### **1. INTRODUCTION**

The development of trade prices and trade flows is important for the performance of small open economies. This is because changes in the price of traded goods affect the terms of trade and thereby the trade balance. Furthermore, the trade elasticities on a bilateral and long-run basis are relevant to designing trade policies and studying international linkages. Another implication of bilateral trade elasticities rests on the size of import and export demand elasticities and whether their absolute values add up to more than unity, a condition known as the Marshall-Lerner (M-L, henceforth) condition in the international trade literature. If it is the case, a depreciation of a home country's currency results in an improvement in the country's external trade balance.

Although, these elasticities also play an important role in predicting how the direction of international trade responds to changes in income and relative prices, they have received very little attention in the empirical literature. In addition to being scant and outdated, the empirical evidence on trade elasticities rests on non-stationary and

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\* We would like to thank Peter Pedroni for generously providing his program procedures that could easily be modified to make the estimations of this study regarding tests for panel unit roots and panel cointegration tests. Valuable comments of one anonymous referee are also greatly appreciated. Of course, we alone remain responsible for any errors.

aggregated data.<sup>1</sup> Examples are Kreinin (1967) Houthakker and Magee (1969), Khan (1974), and Marquez (1990). All these studies suffer not only from the aggregation bias (the exception is Marquez who estimates the trade elasticities on a bilateral basis) but also from the spurious regression problem. The application of conventional econometric techniques to non-stationary (integrated) time series can give rise to misleading results and erroneous inferences.

This study attempts to estimate the long-run bilateral trade elasticities between Sweden vis-à-vis six of its largest trading partners (Denmark, France, Germany, Norway, the UK, and the US) and for the period 1960-1999. The reason for the choice of this period is availability of the data used in this study. The methodology used here is the asymptotic theory of panel cointegration developed by Pedroni (1995, 1997, 1999). This study is the first attempt to test the Swedish bilateral trade elasticities using panel unit root and panel cointegration techniques. One of the major advantages of using a panel cointegration test is a significant increase in power when the cross-sectional dimension of the panel is expanded as compared to the well-known low power of standard cointegration test for small samples. Furthermore, the panel model allows for straightforward panel tests of model specification.

The organization of this paper is as follows: Section 2 introduces the empirical specifications. The estimated results are presented and interpreted in Section 3. Section 4 offers conclusions and policy implications. The methodology for panel unit roots and panel cointegration is described in the Appendix.

## 2. EMPIRICAL SPECIFICATIONS

Following the usual practice in the applied literature on the bilateral import and demand functions, we define the subsequent long-run specification in log-linear and panel form as:

$$\ln IMP_{i,t}^{SWE} = \mathbf{a}_i + \mathbf{b}_i \ln Y_{SWE,t} + \mathbf{d}_i \ln EXR_{i,t} + e_{i,t}, \quad \text{for } i = 1, \dots, N \text{ and } t = 1, \dots, T, \quad (1)$$

$$\ln EXP_{i,t}^{SWE} = \mathbf{a}'_i + \mathbf{b}'_i \ln Y_{i,t} + \mathbf{d}'_i \ln EXR_{i,t} + e'_{i,t}, \quad \text{for } i = 1, \dots, N \text{ and } t = 1, \dots, T, \quad (2)$$

where  $IMP_i^{SWE}$  = the Swedish real import from country  $i$  (Denmark, France, Germany, Norway, the UK, and the US);  $Y_{SWE}$  = the Swedish real income;  $EXR_i$  = real bilateral exchange rate between Sweden and country  $i$  defined as the number of SEK per foreign currency;  $EXP_i^{SWE}$  = the Swedish export to country  $i$  (country  $i$ 's imports from Sweden);  $Y_i$  = the country  $i$ 's real income. (See Appendix 1 for the definitions and sources of the data).  $N$  is the dimension of the panel and  $T$  signifies the time series dimension. In equation (1), if real depreciation is to reduce the Swedish imports, it is

<sup>1</sup> One exception is Bahmani-Oskooee and Brooks (1999) who have estimated the trade elasticities on a bilateral basis by applying Johansen's cointegration method to the US data.

expected that the estimate of  $\delta$  is negative and if an increase in the Swedish income leads to an increase in the Swedish imports, the estimate of  $\beta$  should be positive. In equation (2), if real depreciation of the SEK, i.e., an increase in exchange rate, is to promote the Swedish competitiveness and thus her exports, the estimate of  $d'$  should be positive. Finally, if the income elasticity of  $b'$  is positive, this implies a rise in the Swedish exports to a given country due to the country's economic growth.<sup>2</sup>

The question of whether there is a long-run stable relation between three variables in the panels defined by equations (1) and (2) could be examined by panel cointegration analysis. One of the major advantages of using a panel cointegration test is a significant increase in power when the cross-sectional dimension of the panel is expanded as compared to the well-known low power of standard cointegration test for small samples. The panel model also allows for straightforward panel tests of model specification.

In order to see if the variables are cointegrated in a panel perspective (i.e., if there exists any long-run connection between the variables in the panel) we test first for the integration order of each variable in the panel. A variable is considered to integrate to order  $d$ , denoted  $I(d)$ , if it must be differenced  $d$  times to attain stationarity. For this purpose, we will make use of the tests provided by Levin and Lin (1993), Levin et. al (2002), and Im et. al (2003). If the variables appear to be integrated, we will utilize several test statistics introduced by Pedroni (1995; 1997; 1999). These test methods allow different individual effects across  $N$  or cross-sectional interdependency and take into account the off-diagonal terms in the residual long-run covariance and the effects of spurious regression in the heterogeneous panel. Pedroni argues different types of tests are appropriate for testing the null of cointegration in panel models with heterogeneous dynamics, fixed effects and heterogeneous cointegrating slope of coefficients.<sup>3</sup>

### 3. ESTIMATION RESULTS

The method conferred previously consists of two steps. The first step is to test the variables for stationarity by applying panel unit root tests. From Table 1, we can conclude that each variable for each country in the panel is integrated of the first order;  $I(1)$ . All three tests for panel unit roots give the same results. Thus, it is of paramount importance to test for cointegration in this case, because if the variables do not cointegrate there is not any long-run relationship between them and any estimated results based on the variables in the level form will be spurious. Since the variables are found to be non-stationary of the first order, we investigate whether the variables in each model establish any cointegrating long-run relationship. Based on Pedroni's tests for panel cointegration, presented in Table 2, we find strong empirical evidence for panel cointegration for the export function because all four tests reject the null

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<sup>2</sup> We use the real bilateral exchange rate as a measure of relative prices since import and export prices on a bilateral basis are not available. The real exchange rate is also used by other authors as a measure of relative prices (e.g., Dornbusch 1980; Bahmani-Oskooee and Brooks 1999).

<sup>3</sup> For more details on these methods, see Appendix 2.

hypothesis of no panel cointegration at the one percent significance level. The evidence for panel cointegration is, however, less strong for the import function. One test rejects the null of no panel cointegration at the one percent significance level but the other three tests reject the null only at the ten percent significance level. It should be mentioned however that the estimated value for each test statistic is very close to its critical value at the five percent significance level. Thus, we consider the import function to be a cointegrated panel.

The estimated bilateral elasticities for each function, which are estimated by utilizing the dummy least square method, are presented in Table 3. We can see that the export function is very foreign income elastic, since the bilateral income elasticities are higher than one for all countries in the sample. The bilateral price elasticities are, however, much lower but they all are of the expected sign. The import function as well seems to be very elastic for the domestic income. The bilateral price elasticities are of the expected sign except for France. These elasticities are of lower size (one exception is the case of Germany which is close to one in absolute terms). The sum of bilateral price elasticities of export and import for each country does not add up to more than one in absolute terms except for Germany. Notice that the French case is not considered since the import elasticity is not of the expected sign. However, we can conclude that the M-L condition is not fulfilled except for Germany.

#### **4. CONCLUSIONS AND POLICY IMPLICATIONS**

This study applies the new developments in the field of panel cointegration analysis to investigate the long-run bilateral trade elasticities between Sweden and her six major trading partners for the period 1960-1999. Several tests for panel unit roots and panel cointegration are conducted. The results show that each variable is characterized by one panel unit root. However, the tests for panel cointegration show that both export and import functions can be considered as cointegrated panel systems. The estimated long-run elasticities reveal that the export function is very foreign income elastic but less-price elastic. The same results holds for the import function. The Swedish import function is very elastic to domestic income level but less elastic to terms of trade in real terms except for Germany.

What are the policy implications of our findings? The elasticity approach considers the real exchange rate and its effect on the demand and supply of traded goods as the key factor, while the absorption-approach asserts that total expenditure is the most critical factor for understanding and correcting import and export functions. The dynamics of the trade balance are explained by agents' response to transitory and permanent shocks, in particular, shocks in productivity.

What does cointegration, or lack of cointegration between real trade flows and real exchange rates or real incomes tell us about the state of the economy? Theory asserts that cointegration is to be expected under the maintained hypothesis that the economy is working properly. In a well-functioning economy, without permanent one-sided productivity shocks, cointegration is to be expected. This implies that the lack of

cointegration is the outcome of distorted markets, fundamental policy problems, and the existence of permanent technological shocks to the domestic economy.

However, our results offer three policy implications. First, our findings of cointegration between the variables indicate short-run imbalances are temporary and are sustainable in the long run. Second, macroeconomic policies (such as monetary policies) in Sweden have been less effective. Third, the sums of price elasticities are less than one in absolute terms except for Germany. This implies the Marshall-Lerner condition is not fulfilled except for Germany. Thus, bilateral devaluations are not likely to improve Swedish trade balances except for the bilateral trade with Germany.

## APPENDIX 1

### Data Definition and Sources

All variables are yearly over 1960-1999 period and obtained from the following sources:

- (a) National Board of Trade, Stockholm.
- (b) International Financial Statistics, various issues.

### Variables:

$IMP_i^{SWE}$  = The Swedish real imports from country  $i$  (Denmark, France, Germany, Norway, the UK and the US): nominal import values from source (a) deflated by the Swedish import price index from source (b).

$EXP_i^{SWE}$  = The Swedish real exports to country  $i$ : nominal export values from source (a) deflated by the Swedish export price index from source (b).

$Y_{SWE}$  = Real GDP in Sweden, which is set in index form to make it unit free (source b).

$Y_i$  = Real GDP in country  $i$ , which is set in index form to make it unit free (source b).

$EXR_i$  = Real bilateral exchange rate between Sweden and country  $i$ . It is defined as  $(P_i \cdot E_i / P_{SWE})$ , where  $P_i$  is country  $i$ 's GDP deflator (source b);  $E_i$  is the nominal bilateral exchange rate defined as the number of SEK per currency  $i$  (source b); and  $P_{SWE}$  is the Swedish GDP deflator (source b).

## APPENDIX 2

### Panel Unit Roots and Panel Cointegration

It is well known in the literature that the data generating process for many economic variables are characterized by stochastic trends that might result in spurious inference if the time series properties are not carefully investigated. One of the well-known test statistics for this purpose is the augmented Dickey-Fuller unit-root test. This test in simple form is the following:

$$x_t = \gamma x_{t-1} + v_t, \quad (A1)^4$$

which under the null hypothesis of no unit root, i.e.  $\gamma = 1$ , is equivalent to the following:

$$Dx_t = v_t. \quad (A2)$$

On the other hand, Shiller and Perron (1985) found that the power of Dickey-Fuller unit-root test is very low in small sample sizes. To increase the power of the test, Levin and Lin (1993) and Im, Pesaran, and Shin (2003) (IPS hereafter) suggested panel versions of the test. A panel version of the Dickey-Fuller unit-root test is the following:

$$\begin{bmatrix} x_{1t} \\ x_{2t} \\ \vdots \\ x_{Nt} \end{bmatrix} = \begin{bmatrix} \gamma_1 x_{1t-1} \\ \gamma_2 x_{2t-1} \\ \vdots \\ \gamma_N x_{Nt-1} \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \\ \vdots \\ v_{Nt} \end{bmatrix}, \quad (A3)$$

where N signifies the number of cross-sections. The error terms are assumed to be white noise processes. The null hypothesis of panel unit root is  $\gamma_i = 1$  for all i. The panel unit root test that Levin and Lin (1993) (LL) suggested is based on the following regression:<sup>5</sup>

$$x_{it} = \gamma_i x_{it-1} + v_{it}, \text{ for } i = 1, \dots, N \text{ and } t = 1, \dots, T. \quad (A4)$$

The panel estimator can be defined as the following according to the authors:

<sup>4</sup> Notice that  $x_t$  is a scalar variable.

<sup>5</sup> It should be pointed out that it is possible to add individual constant and trend terms in equation (A3).

$$\sqrt{NT}(\hat{g} - 1) = \frac{\frac{1}{\sqrt{N}} \sum_{i=1}^N \frac{1}{T} \sum_{t=1}^T x_{it-1} v_{it}}{\frac{1}{N} \sum_{i=1}^N \frac{1}{T^2} \sum_{t=1}^T x_{it-1}^2} . \tag{A5}$$

The following t-statistics can be used to test for the null hypothesis of panel unit root:

$$t_g = \frac{(\hat{g} - 1) \sqrt{\sum_{i=1}^N \sum_{t=1}^T x_{it-1}}}{\sqrt{\frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T v_{it}^2}} . \tag{A6}$$

The alternative hypothesis in the LL test is  $g_i = g < 1$  for all  $i$ . The Monte Carlo simulations conducted by Levin et. al (2002) show that the power of the panel-based unit root test is much higher compared to individual unit root tests.

The IPS test allows for a heterogeneous coefficient of unit root and they suggest an average of the individual Dickey-Fuller tests. Their test, which has better size properties, is defined below:

$$\bar{t} = \frac{1}{N} \sum_{i=1}^N t_i , \tag{A7}$$

here  $t_i$  is the individual t-statistic for testing  $H_0: \gamma_i = 1 \forall i, i = 1, \dots, N$ . The alternative hypothesis in the IPS test is  $g_i < 1$  for all  $i$ . That is, it allows for heterogeneity in the panel. Monte Carlo simulations conducted by Karlsson and Löthgren (2000) shows the better performance of the IPS test regarding power properties. Performing unit root tests in a panel perspective is important in order to avoid spurious regression in panel data.

If the variables contain unit roots, a natural next step is to test for cointegration. Pedroni (1995; 1997; 1999) suggests the following test statistics to test for panel cointegration:

1. Panel t-Statistic (Non-Parametric):

$$Z_{tN,T} = \left( \tilde{s}_{N,T}^2 \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \left( \hat{e}_{i,t-1} \mathbf{D} \hat{e}_{i,t} - \hat{\mathbf{I}}_i \right), \tag{A8}^6$$

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<sup>6</sup> Notice the  $e_{i,t}$  represents the error term in the panel equation that is tested for panel cointegration. See equations (1) and (2).

## 2. Panel t-Statistic (Parametric):

$$Z_{iN,T}^* = \left( \tilde{s}_{N,T}^{*2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^* , \quad (\text{A9})$$

## 3. Group t-Statistic (Non-Parametric):

$$N^{-1/2} \tilde{Z}_{iN,T} = N^{-1/2} \sum_{i=1}^N \left( \hat{s}_i^2 \sum_{t=1}^T \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{t=1}^T \left( \hat{e}_{i,t-1} \mathbf{D} \hat{e}_{i,t} - \hat{I}_i \right), \quad (\text{A10})$$

## 4. Group t-Statistic (Parametric):

$$N^{-1/2} \tilde{Z}_{iN,T}^* = N^{-1/2} \sum_{i=1}^N \left( \sum_{t=1}^T \tilde{s}_i^{*2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{t=1}^T \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^* , \quad (\text{A11})$$

where

$$\hat{I}_i = \frac{1}{T} \sum_{s=1}^{k_i} \left( 1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \hat{\mathbf{m}}_{i,t} \hat{\mathbf{m}}_{i,t-s} , \quad (\text{A12})$$

$$\hat{s}_i^2 = \frac{1}{T} \sum_{t=1}^T \hat{\mathbf{m}}_{i,t}^2 , \quad \hat{\mathbf{s}}_i^2 = \hat{s}_i^2 + 2\hat{I}_i , \quad (\text{A13})$$

$$\tilde{s}_{NT}^2 = \frac{1}{T} \sum_{t=1}^T \hat{L}_{11}^2 \hat{\mathbf{s}}_i^2 , \quad \hat{s}_i^{*2} = \frac{1}{T} \sum_{t=1}^T \hat{\mathbf{m}}_{i,t}^{*2} , \quad (\text{A14})$$

and

$$\hat{L}_{11}^2 = \frac{1}{T} \sum_{t=1}^{k_i} \hat{\mathbf{h}}_{i,t}^2 + \frac{2}{T} \sum_{T=1}^T \left( 1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \hat{\mathbf{h}}_{i,t} \hat{\mathbf{h}}_{i,t-s} . \quad (\text{A15})$$

The residual terms used to estimate the above expressions are calculated by running the following regressions:

$$\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \hat{\mathbf{m}}_{i,t} , \quad (\text{A16})$$

$$\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \sum_{k=1}^{K_i} \hat{\mathbf{g}}_{i,k} \Delta \hat{e}_{i,t-k} + \hat{\mathbf{m}}_{i,t}^* , \quad (\text{A17})$$

and

$$\Delta y_{i,t} = \sum_{m=1}^M b_{mi} \Delta x_{mi,t} + \hat{\mathbf{h}}_{i,t} . \quad (\text{A18})$$

Pedroni provides some adjustments for each of all test statistics (both for panel unit root tests and panel cointegration tests) described above that result in standard normal distributions. In this study, we report the adjusted values so that in all cases the reported values should be compared to the  $N(0,1)$  distribution. This is true for both the cointegration and unit root tests.<sup>7</sup>

**Table 1. Test Results for Panel Unit Roots**

H <sub>0</sub> : I(1), H <sub>1</sub> : I(0)			H <sub>0</sub> : I(2), H <sub>1</sub> : I(1)				
	LL <sub>1</sub>	LL <sub>2</sub>	IPS		LL <sub>1</sub>	LL <sub>2</sub>	IPS
$\ln IMP_i^{SWE}$	0.88	1.14	0.80	$\ln IMP_i^{SWE}$	-8.52 <sup>a</sup>	-7.45 <sup>a</sup>	-11.0 <sup>a</sup>
$\ln EXP_i^{SWE}$	0.09	-0.14	-0.10	$\ln EXP_i^{SWE}$	-9.58 <sup>a</sup>	-8.38 <sup>a</sup>	-11.06 <sup>a</sup>
$\ln Y_{SWE}$	0.60	1.74	0.39	$\ln Y_{SWE}$	-5.01 <sup>a</sup>	-4.89 <sup>a</sup>	-6.47 <sup>a</sup>
$\ln EXR_i$	0.41	0.28	-0.61	$\ln EXR_i$	-8.18 <sup>a</sup>	-8.26 <sup>a</sup>	-11.32 <sup>a</sup>
$\ln Y_i$	0.43	0.22	0.29	$\ln Y_i$	-6.53 <sup>a</sup>	-6.12 <sup>a</sup>	-8.11 <sup>a</sup>

Notes: LL<sub>1</sub> and LL<sub>2</sub> are the tests suggested by Levin and Lin (1993). The first test augments the regression until autocorrelation is removed. The second test takes into account the effect of potential autocorrelation when the parameters are estimated. IPS is the test suggested by Im et. al (2003). a signifies that the null hypothesis can be rejected at 1% significance level. The adjusted test results are presented here which can be compared to the  $N(0,1)$  distribution. Notice that each test is one sided (to the left side of the distribution).

**Table 2. Panel Cointegration Test Results for Import and Export Functions Based on Pedroni tests**

	Test 1	Test 2	Test 3	Test 4
Export Function	-2.50	-2.19	-2.25	-2.31
Import Function	-1.57	-2.00	-1.50	-2.56

Notes: Notice that Test 1 = Panel t-Statistic (Non-Parametric), Test 2 = Panel t-Statistic (Parametric), Test 3 = Group t-Statistic (Non-Parametric), and Test 4 = Group t-Statistic (Parametric) as described in the main text. Once again using Pedroni’s procedure, we present the adjusted values here that can be compared to the  $N(0,1)$ . Since the tests

<sup>7</sup> For more details see Pedroni (1999). These methods are also presented in Baltagi (2001).

are one sided the 1% critical value is  $-1.96$ , the 5% value is  $-1.64$  and the 10% critical value is  $-1.28$ .

**Table 3. The Long-Run Bilateral Trade Elasticities**

	Export Function		Import Function	
	$REX_i$	$Y_i$	$REX_i$	$Y_{SWE}$
Denmark	0.03	1.33	-0.74	
France	0.26	1.21	0.79	
Germany	0.59	1.29	-0.91	
Norway	0.07	1.46	-0.16	
UK	0.16	1.40	-0.05	
US	0.24	1.35	-0.51	
Sweden				1.42

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## **PUBLIC OPINIONS OF THE IMPEACHMENT OF PRESIDENT WILLIAM JEFFERSON CLINTON: A LOOK BACK**

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### **ABSTRACT**

*The purpose of this paper is to determine how gender, party affiliation, political views, age, race, education, income, attendance at religious services, and other such variables affected public opinions of President Clinton shortly after the release of the Starr report (and before the impeachment vote in the House of Representatives). Using data from a CBS NEWS / NEW YORK TIMES poll taken from September 12 to September 15, 1998, this study found the following: 1) Age had an effect on opinions concerning the Clinton scandal with younger individuals more likely to want the impeachment process to begin. This information provides some support for the Life-experience hypothesis. 2) Women were more likely to want Clinton to resign and less likely to want impeachment, which is perhaps consistent with the strong support Clinton had from women voters in the 1996 election. 3) Those with more education and income were more likely to want Clinton to resign and less likely to want impeachment, which is consistent with the resource hypothesis. 4) The combination of all attributes and individual characteristics determine how final opinions are established concerning resignation, impeachment, and dropping the matter.*

**JEL Classification Codes:** D72

**Key words:** *Public Opinion; impeachment; President William Jefferson Clinton*

### **1. INTRODUCTION**

On January 17, 1998, President William Jefferson Clinton testified in the Paula Jones trial and denied having sexual relations with a White House intern, Monica Lewinsky. On August 17 of the same year, the president again testified, this time before Kenneth Starr's grand jury. That night President Clinton addressed the nation and stated that he had been involved in an inappropriate affair.

On September 11, 1998, Kenneth Starr and the Office of the Independent Council submitted a 455-page report to the House Judiciary Committee, outlining eleven impeachable offenses that the president was being accused of. The Judiciary Committee which consisted of twenty-one Republicans and sixteen Democrats, reduced and revised these to four allegations that were then voted on: article one "alleges that on Aug. 17, 1998, William Jefferson Clinton willfully provided perjurious, false and misleading testimony before Independent Counsel Kenneth Starr's grand jury"; article two "alleges that the president willfully provided perjurious, false and misleading