Akdeniz İ.İ.B.F. Dergisi (13) 2007, 185-199

MINIMUM WAGE AND AVERAGE WAGE RELATIONSHIP IN TURKEY: A COINTEGRATION AND ERROR CORRECTION ANALYSIS

TÜRKİYE'DE ASGARİ ÜCRET VE ORTALAMA ÜCRET İLİŞKİSİ: BİR KOENTEGRASYON VE HATA DÜZELTME ANALİZİ

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ABSTRACT

The effect of statutory minimum wages on average wages is not a well-studied subject in Turkey. Using cointegration analysis this study shows the existence of a long-run relationship between real minimum wages and real private-sector average wages. The short-run adjustment properties of the relationship were also established via error-correction model. The Granger-causality relationship was found to be bi-directional. The estimated results imply that the minimum wages have considerable potential to influence major macroeconomic variables in Turkey.

Keywords: *Minimum Wages, Average Wages, Cointegration.*

ÖZET

Yasal asgari ücretlerin iş piyasasındaki özel-sektör ortalama ücretlerle olan ilişkisi Türkiyedeki araştırmalarda pek üzerinde durulan bir konu olmamıştır. Bu çalışma, eşbütünleme yöntemini kullanarak asgari ücretlerle özel-sektör ortalama ücretleri arasında uzun-dönem bir ilişkinin olduğunu ortaya koymaktadır. Söz konusu ilişkinin kısa dönemde gösterdiği özellikler ise 'hata düzeltme' modeli kullanılarak analiz edilmiştir. Sonuçlar değişkenler arasında çift yönlü işleyen bir Granger-nedenselliğinin bulunduğunu göstermektedir. Elde edilen bulgulara göre, asgari ücretlerin özellikle özel sektördeki ortalama ücretler vasıtasıyla ekonomideki makro deşkenleri önemli ölçüde etkileme potansiyeli taşıdığını söylemek mümkündür.

Anahtar Sözcükler: *Asgari Ücretler, Ortalama Ücretler, Koentegrasyon.*

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INTRODUCTION

Through minimum wage laws the government makes it illegal for employers to pay, and workers to work for, a wage that is less than a fixed minimum amount. Economic theory, however, states that statutory minimum wages raise unemployment, which is already an important labor market issue especially in developing countries. According to the theory, in the absence of minimum wage laws, though lowest wages determined by market forces will be lower than the enforced minimum, market employment level will be higher than the employment level associated with the minimum. Many empirical studies, such as Keil et al. (2001), Kramarz and Philippon (2001), Neumark et al. (2000), Partridge and Partridge (1999), found significant negative relationship between minimum-wage enforcement and employment level.

However, this traditional theory and supporting empirical literature have been recently challenged by oppositionist view. Some of the representative studies are Krueger (2001), Card and Krueger (2000, 1995, 1994), Dickens et al. (1999), Lang and Kahn (1998), Machin and Manning (1997), and Bell and Wright (1996). The challengers argue that it is not possible to adequately determine employment effects of statutory minimum wages by just considering demand and supply components. Other than this, they suggest, human capital investment rates, presence of employer's monopsony power, matching efficacy between workers and vacancies, segmentation in labor markets, and other limiting factors should also be taken into consideration.

Most of the studies conducted by oppositionists have been rigorously criticized in various grounds.[1] The findings of Card and Krueger (1994) were the primary source of controversy, as they concluded that minimum wages do not increase unemployment. Though this finding apparently conflicts with basic economic theory, various subsequent research, as cited above, continued to indicate that in the minimum wage case the estimated labor demand curve may not be that much elastic. Nevertheless, the dispute over the possible employment effects of minimum wages continues.

An important reason for minimum wages to attract so much controversy is clearly due to the presence of various estimation difficulties. As Stock and Watson (2003) state, since "prices and quantities are determined by supply and demand, the OLS estimator in a regression of employment against wages has simultaneous causality bias" (p. 403). Many estimates of minimum wages are prone to misspecification and are subject to incorrect estimation approach. With this regard, Mills et al. (1999), for example, could not find any significant effect using natural experiment

[1] See, for example, Neumark (2001), Burkhauser et al. (2000), Abowd et al. (1999).

approach, but using time series analysis they found a significant and negative relationship between minimum wages and employment. Other researchers, such as Bazen and Marimoutou (2002), Deere et al. (1995), also reported that the estimates are sensitive to the specification adopted.

Considering the problematic nature of direct employment elasticity estimates, it may be wise to apply advanced econometric techniques to see if there is a meaningful relationship between statutory minimum wages and average wages. In this way, it could be possible to determine, implicitly in some cases, whether a minimum-wage hike distorts other labor market variables, or its impact is limited to its first-hand domain. This approach might be even more advantageous in developing countries where detailed data is not available to apply a correctly specified procedure. Despite such possible advantages, in the literature less attention has been paid to the relationship between minimum and average wages. To the best of our knowledge, there is no domestically conducted research in this subject.

The evidence on the impact of statutory minimum wages on the other wages is mixed, albeit the majority shows significant positive effect. Using the Trinidad and Tobago labor force survey Strobl and Walsh (2003) show that the introduction of a national minimum wage decreased the probability that a low-wage worker earns less than the minimum wage by about 19%. They detected differential impacts across firm size implying that as the firm size increases, probability of earning above the minimum also increases. Maloney and Nunez (2002) use panel data from Colombia and identify that the minimum wage can have a significant effect on the public sector wage distribution at the near-minimum (i.e. 10% up or down) wage levels. More specifically, they find that a 1% increase in the minimum wages raises near-minimum wages of public sector employees by about 0.6%. Fajnzylber (2001) finds even larger wage effects for Brazil. He uses longitudinal data from Brazil's Monthly Employment Survey over the 1982-97 periods. His estimates reveal that a 1% increase in the minimum wages raises the wages of unregistered paid employees and self-employed workers by 1.03% and 1.33%, respectively. For the public sector workers the elasticity was 1.08. These imply that the minimum wage elasticity of wages is positive and greater than one. Fajnzylber argues that these findings reflect the tendency that especially private sector workers try to adjust their wages by looking at the minimum wage changes. Without making any distinction between public and private sectors, Rama (2001) analyzes the effects of doubling the minimum wage, in real terms, on wage earnings. He uses data from the 1993 Indonesian labor force survey. His results show that the 100% minimum wage hike increased average wages by only 10 percent. Bell (1997) uses comparative data from Mexico and Colombia to determine the effect of minimum wages. She finds that minimum wages have no effect on wages. She states that this finding is the consequence of very low, only being about

13% of the average unskilled employee wage, minimum wages in Mexico. In Colombia where the minimum wages being set closer to the average wages, however, the results reveal that minimum wages have a significant impact on wages, having an elasticity of 0.37. Grossman (1983) investigates how changes in the minimum wage affect other wages. He concludes that the legal minimum wage causes wage increases for employees slightly above the minimum. Grossman postulates that this may happen for two reasons. First, workers especially just above the minimum will not want to lose their relatively higher wage status, and thus bargain accordingly with employers. Employers also avoid from severe relative wage settings due to possible adverse effects on work effort. Second, the increase in the minimum increases the demand for above minimum wage employees, as the better becomes the cheaper. In an earlier study, Gramlich (1976) also studies the impact of minimum wages on other wages. He finds that raising the minimum wage raises average wages causing higher compliance costs.

The main objective of this paper is to determine the possible existence of a long-run and short-run relationship between legal minimum wages and average private-sector wages. In the minimum and public sector wage settlements, in which the government is the dominant actor, one of the first objectives of Turkish governments almost always has been to keep the range unchanged or to increase minimum wages slightly above the increase in the public sector wages. Therefore, in this study we ignore the effect of legal minimum wages on public-sector wages. In section two of the paper we briefly introduce the state of minimum wages and the data used. Section three describes the methods of analysis and presents the estimated results. Finally, section four provides concluding remarks.

MINIMUM WAGES AND DATA

The regulation of minimum wages in Turkey dates back to early 1970s. The first practice of statutory minimum wages was made in 26 cities, divided in six groups, covering the years 1969-1972. Including the big ones like İstanbul, Ankara and İzmir, most of these cities were paid-worker intensive. Between the years 1972-1974 minimum wages were set in 59 cities. After mid-1974, the minimum wage practice covered all 67 cities. According to the Labor Law (1475/33, amended by 4857/39 in 2003) the Minimum Wage Fixing Committee must determine the statutory minimum wages at least in every two years. The committee consists of equal number of employee, employer and government representatives and decides by majority vote. The minimums are set separately for under 16 years olds. Until 1989 the minimum wages were determined separately for industry-services and agriculture-forestry workers, but now only one legal minimum is set.

The labor legislations and even the Turkish Constitution (article 55) states that in determining the legal minimum wages, minimum living standards of the worker and economic condition of the country must be taken into consideration. The expenses necessary to satisfy minimum living standards include expenditures made for food, housing, health and recreation expenses, commuting and other essentials. Nevertheless, minimum wages always stayed under the minimum living-standard expenses in Turkey. The Minimum Wage Fixing Committee's primary consideration, due to the government and employer representatives, mostly has been the economic condition of the country rather than the level of minimum living standards. From the first application of the legal minimum wages to today, the minimum wages remained to be around 20 percent of the average wages.

The cost of minimum wages has always been very much higher than the net minimum wages in Turkey. In recent years the net minimum wages are somewhat higher than 50% of cost of the minimum wages. Hence, employers repeatedly complain that they bear the heavy burden of legal minimum wages. According to the employers, the load is more than bearable levels and thus distorts major macroeconomic variables and encourages underground economic activities. The employees, on the other hand, believe that the minimum wages are no way close to the minimum expenses to assure minimum living standards. As a third party, the governments also have been reluctant to undertake the burden due to budgetary concerns. The minimum wage earners must pay 15% income tax. They also pay for other cuts amounting around 15%. Under these conditions, it is essential that the government deals with some of the burden. At least, the government can try to waive the income tax collected from minimum wage earners.

Our annual data for private-sector average wages (W) and legal average minimum wages (MW) begins from 1973 and ends in 2004. The W and MW are, rather than being net, in gross values, measured in Turkish Lira, and the minimum wages are for the age of 16-years and over. Both series are converted into real terms (1963 = 100) via Wholesale Price Indices (WPI), reported by Istanbul Chamber of Commerce. The choice of WPI is intentional since, for this sample period, the use of other indices is rather destabilizing and less convenient, i.e. in unit root testing. The W is constructed from State Institute of Statistics and State Planning Organization electronic data resources. The MW is obtained from Ministry of Labor and Social Security electronic data releases. Bulutay (1998: 169-175) is also a good source especially for the minimum wage data. As mentioned above, the minimum wage data in 1973 covered only 59 cities, but we do not see this as a problem to adversely affect our estimates.

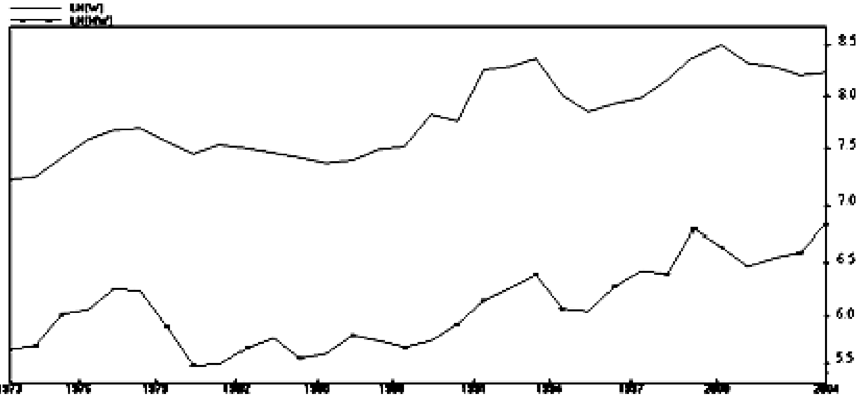

Figure 1: Real Average Wages (W) and Real Minimum Wages (MW) (In Log Form)

Figure 1 above shows the behavior of the private-sector average real wages and real minimum wages between 1973 and 2004. It is clear that both series were adversely affected from the major macroeconomic disturbances observed in Turkish economy in the past, such as in late 1970s, in 1994, and in 2001. Since the vertical line gives the natural log values, to obtain the real raw values of the series, obviously, one jut needs to consider the anti-log of the given numbers, i.e. 5.5 corresponds to about 244.692, and 8.5 to 4914.767 etc.

METHODOLOGY AND ESTIMATION

The cointegration analysis examines whether two or more time-series variables share a common stochastic trend. Using cointegration approach, this paper investigates the long run relationship between statutory minimum wages and average private-sector wages. Let's assume that we have the following two time series in the form,

$$W_t = \alpha_0 + \alpha_1 MW_t + \varepsilon_t \qquad (1)$$

where W_t is the average private sector wages in period *t*, MW_t is the legal minimum wage in period *t*, ε_t is the error term and α_0, α_1 are parameters. Cointegration requires that all the series in the relationship should have the same order of integration. The order of integration is the number of times, i.e. I(d), that a variable has to be differenced to become stationary. The series W_t and MW_t are said to be cointegrated if they are each I(1) but there exists a linear combination of them, $\varepsilon_t = W_t - \alpha_0 - \alpha_1 MW_t$, that is I(0). In this case, α_1 is called the cointegrating parameter and equation (1) is the cointegrating regression. When W_t and MW_t are cointegrated, least squares estimation of equation (1) provides a super consistent estimator of α_1. This indicates the long run, steady-state equilibrium relationship between W_t and MW_t (Griffits et al., 1993: 701). If the series are not cointegrated of the same order and are nonstationary, then the OLS regression in equation (1) breaks down.

Since many macroeconomic series appear to be nonstationary, we first need to check for the stationarity of the series in equation (1). Several tests exist to check for stationarity of the series. We apply Augmented Dickey-Fuller (ADF) (1981) and Phillips-Perron (PP) (1988) tests. The ADF and PP tests are based on the ordinary least squares regression of the following specifications (2) and (3), respectively:

$$y_t = \eta + \alpha t + \rho y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \tag{2}$$

$$y_t = \eta + \alpha t + \rho y_{t-1} + \varepsilon_t \tag{3}$$

where η, α, ρ and γ are parameters, y_t is the individual series of interest, Δ is the first difference operator, t is a linear time trend and k is determined by Akaike's information criterion to ensure that ε_t is white noise. In order to test whether y_t contains a unit root (i.e., H_0: $\rho = 1$) the OLS estimator for ρ is obtained and for the null related statistics are computed. Davidson and MacKinnon (1993) finite sample critical values are used to determine statistical significances for both ADF and PP tests.

The results of the ADF and PP unit root tests for levels and first-differences are presented in Table 1.

Table 1: Results from the Unit Root Tests

		Augmented Dickey-Fuller (ADF)		Phillis-Perron (PP)	
Variables		Constant, No Trend	Constant, Trend	Constant, No Trend	Constant, Trend
W	Level	-1.3976	-2.1666	-1.4839	-2.4273
	First Diff.	-3.7810*(1)	-4.6429*(1)	-5.0903*	-5.1174*
MW	Level	-1.5141	-2.2348	-1.5989	-2.4310
	First Diff.	-3.4356**(2)	-3.6155**(2)	-5.9011*	-5.9424*

The significance levels are indicated by one asterisk (1%) and two asterisks (5%).
The numbers in parentheses are the lag lengths. The optimal lag lengths are determined by Shazam default. The variables are in log forms.

The ADF and PP unit root test results in Table 1 reveal that the null hypothesis of presence of unit root is not rejected in levels. However, the null hypothesis is rejected for the first differences, implying that W and MW series are stationary in their differences, i.e. I(1). Furthermore, in addition to the stochastic trend, according to the test results, the series exhibit stationarity around the deterministic linear time trend. However, testing trend pattern of the series by utilizing the test equation $z_t = \eta + \beta t + \varphi z_{t-1} + \varepsilon_t$ does not confirm the presence of deterministic trend.

Having confirmed the existence of unit roots for the series, the next step requires testing for cointegration to detect possible long-run equilibrium relationship. If we consider the regression model $y_t = \eta + \rho x_t + \varepsilon_t$, where y_t and x_t are both I(1), one approach to test for cointegration just requires assessing whether the errors in the equation are stationary. This approach is due to Engle and Granger (1987) and demands use of Dickey-Fuller (DF) or Phillips tests on the residuals of cointegrating regressions. In order to decide whether the OLS residuals have a unit root, however, a specially constructed set of tables should be used in determining the critical values for the tests. Table 2 below presents the results of DF and Phillips tests for cointegration.

Table 2: DF and Phillips Test Results for Cointegration

	Dickey-Fuller(DF) test		Phillips test	
	Calculated test statistics	Asymptotic Critical values (10%)	Calculated test statistics	Asymptotic Critical values (10%)
Constant, No trend	-3.3271	-3.04	-2.5623	-3.04
Constant, Trend	-3.6754	-3.50	-2.7720	-3.50

Regressand is W, with H_0: no cointegration. Given the AIC & SC criterions optimal lag length is selected to be 4.

Results from Table 2 indicate that only DF test provides evidence for the existence of cointegration between the W and MW series. The Phillips test does not support this finding. In order to alleviate this ambiguous outcome, it is better to conduct an additional cointegration test. For this purpose we will perform the Johansen-Juselius (JJ) test.

The JJ procedure uses two test statistics to determine the number of cointegrating vectors. They are the trace and the maximum eigenvalue test statistics. In utilizing $\lambda_{trace} = T \sum_{j=r+1,n} ln(1-\lambda_j)$ equation the trace test statistic, for the null, hypothesizes that there are at most r number of cointegrating vectors. In the equation T represents the number of observations, and λ_js show the estimated values of the characteristic roots, in assuming that the series are I(1). Using $\lambda_{max} = -T\, ln(1-\lambda_{r+1})$ relationship the maximum eigenvalue test statistic constructs the null hypothesis as there are at most r

cointegrating vectors, and the alternative hypothesis as *r+1* cointegrating vectors[2].

The results of Johansen-Juselius cointegration tests for the series are given in Table 3 below. Special critical values for the test statistics are obtained from Johansen and Juselius (1990).

Table 3: JJ cointegration test results

	r = 0	Critical values	r <=1	Critical values
λ_{trace}	18.436	17.08 (5%)	0.268	6.7 (10%)
λ_{max}	18.168	14.60 (5%)	0.268	6.7 (10%)

According to the AIC and SC criteria optimum lag length is selected to be 5. No restrictions on the constant term are imposed.

Results presented in Table 3 show, at the 5% level of significance, that there is a cointegration vector indicating that the two series are cointegrated. Considering the DF and JJ cointegration test results together, we can conclude that W and MW are cointegrated. This finding reveals that there is a long-run association between real minimum wages and real average private sector wages. As we determine that the series are cointegrated, we can further proceed in order to observe the short-run properties of the series, using error correction models.

Engle and Granger (Granger, 1983; Engle and Granger, 1987) demonstrate that the cointegrating variables must have an error correction model (ECM) representation. Further, if the series are cointegrated, then the possibility of the estimated regression being spurious is ruled out due to the problems such as autocorrelation, endogeneity and omitted variable bias. The short-run relationship between the series can be estimated using the following error correction model:

$$\Delta \ln W_t = \alpha + \sum\theta(i)\Delta \ln MW_{t-i} + \sum\Phi(i)\Delta \ln W_{t-i} + \psi E_{t-1} + \varepsilon_{1t} \quad (4)$$

$$\Delta \ln MW_t = \gamma + \sum\delta(i)\Delta \ln W_{t-i} + \sum\Omega(i)\Delta \ln MW_{t-i} + \varphi E_{t-1} + \varepsilon_{2t} \quad (5)$$

In equations (4) and (5) Δ is the first-difference operator, ε_{1t}, ε_{2t} are white noise residuals, and $\alpha, \theta, \Phi, \psi, \gamma, \delta, \Omega, \varphi$ are parameters to be estimated. The expression E_{t-1} represents error correction term which is in fact residuals from cointegration regression, i.e., $\ln W_{t-1} - \beta_1 - \beta_2 \ln MW_{t-1}$. In the model, the coefficients of E_{t-1}, ψ and φ, capture the adjustments of $\Delta \ln W_t$ and $\Delta \ln MW_t$ towards long-run equilibrium, and the related coefficients for $\Delta \ln W_{t-i}$ and $\Delta \ln MW_{t-i}$ capture the short-run dynamics of the model.

[2] For more info about these tests see Enders (2004, p 352).

The error correction model estimation results coming from equation (4) and (5) are given in Table 4 below:

Table 4: The Output for the Error Correction Models

Equation (4): Dep. Var. : ΔlnW_t		Equation (5): Dep. Var. : $\Delta lnMW_t$	
Variable	Coefficient	Variable	Coefficient
E_{t-1}	-0.30839**	E_{t-1}	-0.39781***
$\Delta lnMW_t$	0.38700*	ΔlnW_t	0.84068*
ΔlnW_{t-1}	0.01672	$\Delta lnMW_{t-1}$	-0.06636
$\Delta lnMW_{t-1}$	0.04999	ΔlnW_{t-1}	0.09448
ΔlnW_{t-2}	0.13693	$\Delta lnMW_{t-2}$	-0.15379
$\Delta lnMW_{t-2}$	0.07079	ΔlnW_{t-2}	0.01249
Constant	0.01558	Constant	-0.00365

Significance levels are shown as one asterisk (1%), two asterisks (5%), and three asterisks (10%). Durbin-Watson and R^2 statistics for equation (4) and (5) are, respectively, 1.87, 1.88 and 0.4210, 0.4213.

Results from equation (4) shows that a positive exodus of lnW_t from its equilibrium value in the preceding period causes a negative change in ΔlnW_t meaning that lnW_t returns back to its long-run equilibrium value. The speed of this correction is determined by the estimated coefficient of -0.30839. A similar interpretation can be made for equation (5). Consequently, findings from the error correction model reveal that there exists bi-directional causality, in the Granger sense, between the average wages and minimum wages. In essence, the estimated values of ψ and φ give the long-run relationship between the series. The other statistically significant coefficients 0.38700 and 0.84068 should be interpreted as short-run responses. It is interesting to see that the average wages influence minimum wages more drastically then the other way around. This finding divulges that the minimum wages are, in fact, not exogenous in Turkey.

CONCLUSION

The analysis conducted above indicates that the variables W and MW are nonstationary, stationary after differencing once, and are cointegrated. Therefore, the regression in levels gives a consistent estimator of the long-run cointegrating relationship between private-sector average wages and legal minimum wages. In the Granger-sense, causality between the series is bi-directional and the short-run relation appears to be spontaneous, meaning that the short-run effects do not spread to the preceding periods.

The estimated results have some important macroeconomic implications. First, they reveal that an increase in minimum wages cause a rise in average private-sector wage rates. This may be due to a 'rangetracking' effect observed in labor markets. Especially in proximities, lower wage earners always try to catch up higher wage earners, and the higher wage earners try to increase or keep the wage gap. Considering this, the governments in Turkey can shape their decisions accordingly, by looking at the presence of inflationary or deflationary pressures in the economy. Since the adverse effects of average wages on employment is a well documented case, the unemployment effects of the legal minimum wages should be monitored more closely by the governments. The second important conclusion to draw from the results is the presence of causal effect of the average wages on the minimum wages. This conclusion simply means that, concerning the minimum wages, the governments in Turkey make their decision based on market forces. Apparently, this is why the statutory minimum wages always have been much under the cost of minimum living standards. The unemployment pressure pushes the level of minimums down, and the government, obeying the labor market forces, can not set and enforce higher minimum wage rates.

There is no doubt that the statutory minimum wages are already lower than reasonably satisfying levels in Turkey. In order to make the minimum wage earners better off, it seems that the government should undertake some of the burden. The government can try to set its policy regarding this issue to decrease the substantial difference between the cost of minimum wages and the net minimum wages.

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Data Appendix

Year	W (Nominal)	MW (Nominal)	W (Real)	MW (Real)
1973	3167	692	1337,981	292,3532
1974	4126	915	1373,502	304,5939
1975	5426	1350	1621,638	403,4668
1976	7556	1650	1924,605	420,2751
1977	10497	2550	2081,499	505,6514
1978	16433	3825	2120,934	493,6758
1979	25374	4875	1869,86	359,2483
1980	43500	6550	1684,806	253,6891
1981	63163	8850	1823,833	255,544
1982	78096	13100	1770,684	297,0185
1983	95448	18281	1689,554	323,5976
1984	134074	22443	1620,7	271,2932
1985	181872	32962	1551,24	281,1426
1986	237344	49637	1587,224	331,9446
1987	364977	66062	1752,389	317,1879
1988	603108	100125	1800,602	298,927
1989	1326600	175500	2400,404	317,5568
1990	1878000	303750	2271,801	367,4438

1991	4536000	575250	3596,749	456,1353
1992	7791000	1071000	3696,078	508,0862
1993	12995000	1885875	3971,249	576,3201
1994	20653000	3056250	2862,583	423,6077
1995	33323000	5602500	2456,881	413,0683
1996	61689000	12022500	2650,683	516,5886
1997	115528000	24688125	2782,315	594,5756
1998	230493000	40605000	3288,741	579,364
1999	406594000	85837500	4037,662	852,4051
2000	698361000	114300000	4511,905	738,4587
2001	929295000	151612500	3807,094	621,1193
2002	1309758000	236437875	3700,684	668,0486
2003	1506221700	306000000	3438,481	698,5527
2004	1704666409	433575000	3514,813	893,9785