Reexamining the Border Tax Effect: A Case Study of Washington State

by

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Abstract: Without an income tax, Washington State relies heavily upon its sales tax revenue to fund public goods and services. Bordering Idaho and especially Oregon, where the sales tax is substantially lower, the juxtaposition of the different tax structures generates the border tax effect in Washington's border counties. Controlling for unobservable county-specific characteristics and spatial autocorrelation, we find that the price elasticity generated by the sales tax discrepancy over the years 1992 – 2006 is -3.11. We estimate that elimination of the sales tax differential between Washington and its neighboring states would generate tax revenue in excess of $145 million at the state level and over $21 million at the county level in border counties.

I. Introduction

The juxtaposition of contiguous governmental jurisdictions with markedly different taxing systems has long provided opportunities for careful evaluation of these systems and their effect on consumer, firm and governmental welfare. The purpose of this paper is to analyze the effect Washington State's retail sales tax has on border counties within the state. Washington is one of only seven states without an income tax and relies heavily upon its sales tax to generate revenue at both the state and local levels. Washington has the sixth highest state sales tax in the nation at 6.5%, with local options adding an additional 0.5-2.4%, for a total of 7.0-8.9%. Thus counties on the border with Oregon, which has no sales tax, and Idaho, which has a 6.0% sales tax with little to no local tax, provide a unique opportunity to investigate and test the border tax effect.

In 1992 the U.S. Supreme Court ruled, in Quill Corp. v. North Dakota, that a business has to be physically present in a state before that state can collect sales taxes (Rothenburg, 2007). Consequently, state and local governments continually battle for the location of businesses in an effort to expand their tax base, increase revenue, and fund public goods. The large sales tax discrepancy between Washington and Oregon (and to a lesser extent, Idaho) provides an easily obtained opportunity for Washingtonians along the border to evade paying their "fair share" of taxes. This tax evasion is not only illegal, due to Washington's use tax, but also decreases the revenue for the state and the local jurisdictions in which the residents live. Furthermore, it may have an adverse impact on the employment level in these border counties.

The heavy reliance of both Washington and Oregon on one type of taxation (sales and income taxes, respectively), makes the governments of these states particularly sensitive to macroeconomic fluctuations which may limit their ability to generate sufficient revenue during recessionary periods. Washington's reliance on sales tax revenue places a disproportionate burden on the lower and middle income residents of Washington that the upper income classes do not face. The regressive nature of a sales (excise) tax and lack of a progressive income tax, led McIntyre (2000) of the nonprofit, non-partisan Institute of Taxation and Economic Policy, to label Washington's tax system the most regressive in the nation. Given the fiscal constraints of the tax system, Washington lawmakers are faced with the inevitable task of raising the marginal tax rates on the lower and middle classes in times of economic downturn.

The border tax effect in the case of Washington State has generated surprisingly little research over the past two decades. Our review of previous literature uncovered only two studies that have examined the border tax problem in Washington State since the 1980s. Brown (1990) uses 1975-1987 data from 11 border cities in Washington to demonstrate the sizeable effect that differences in relative sales tax rates have on retail sales in border counties. A second study by Beck (1992), uses data from 1984 to 1988 to estimate price elasticity measures in order to evaluate policy changes related to sales tax increases in 1983-84. [1] Using 2006 data, we illustrate in Figure 1, the discrepancy in sales between Washington's border counties and the rest of the state. Figure 2 shows that this pattern is especially pronounced along the Washington-Oregon border, where consumers may choose to purchase retail goods at a significantly lower taxing level.

To the best of our knowledge, no new studies have examined the border tax problem in Washington State since the late 1980s. In light of this, the present study makes several important contributions to the literature. First, we provide fixed effects estimates of the price elasticity

[1] Washington's legislature passed an increase in the state sales tax rate in 1983 to 6.5% while allowing for some border counties (Clark, Cowlitz, Skamania and Klickitat) to maintain their tax rate at 5.4% in order to reduce the border tax effect. In 1984, the Washington State Supreme Court overturned this legislation as the state constitution requires uniform taxes throughout the state. However, Beck's evidence validates the legislature's reasoning to combat the border tax effect.

associated with taxable retail goods in the 39 Washington counties and evaluate the border effect using data from 1992 to 2006. Second, we estimate a spatial lag model to investigate whether the variation in taxable retail sales in our sample is significantly affected by spatial dependence among counties in our sample. Finally, we use the estimated county relative price to generate the potential sales and tax revenue gain that could arise if either partial or complete tax harmonization were to occur between Washington and its neighboring lower-tax states.

Our fixed effects results show that the estimated border tax effect is both statistically and economically significant in Washington State. We estimate the price elasticity of retail goods in Washington at -3.11. This result is in line with estimates of border tax effects by Brown (1990) and Beck (1992) and corresponds to approximately $2.2 billion in lost sales for the state of Washington. This loss of sales translates, on average, to over $145 million in forgone state tax revenue and over $21 million in foregone county tax revenue for the year 2006. Furthermore, our empirical results remain robust to spatial autocorrelation corrections and we note two observations regarding the latter. First, the estimated coefficients in our fixed effects specification are remarkably stable when compared to maximum likelihood estimates obtained from the spatial lag specification. Second, we do not find evidence of spatial dependence in our sample. Overall, these conclusions from our spatial analysis are in line with similar findings in recent research (see for example, Blonigen et al., 2007; Tosun and Skidmore, 2007).

The remainder of this paper is structured as follows: Section II reviews literature related to tax harmonization and empirical estimates of the border tax effect. Section III details the empirical methodology related to our fixed effects and spatial lag models and describes the data. Empirical results are presented in Section IV along with retail sales tax revenue estimates by county. We offer concluding remarks and direction for future research in Section V.

II. Background

A. Theoretical Considerations

Tax harmonization is a major topic in examining the tax structures of neighboring jurisdictions and often prescribed as a means to minimize the border tax effect (see for example Behrens et al, 2007; Conconi et al, 2008). Implementation of full or partial tax harmonization however, is difficult in practice due to the differential impact on tax revenue in the countries (or regions) involved. For example, in a two-country, non-cooperative setting, Kanbur and Keen (1993) and Nielsen (2001) find the tax rate and gross tax revenue to be higher in the large country, while per capita revenue is greater in the small country by strictly undercutting the large country to gain a fiscal advantage[2]. Both studies find the non-cooperative behavior to be Pareto inefficient and tax harmonization to be beneficial for the large country and detrimental to the small country. In place of full tax harmonization, both purport a binding minimum tax (set in between the small and large country tax rates) to increase tax revenues in both countries. These results are in line with Mintz and Tulkens (1986) who find Pareto improving tax rates never reduce taxes in both regions (countries) and always increase in the tax importing region (small country).

The experience of the European Union (EU) is particularly informative with respective to tax harmonization efforts. Specifically, the standard Value Added Tax (VAT) on goods and services in the EU ranges between a minimum of 15% (e.g. Cyprus) to a maximum of 25% (e.g. Denmark). At present, the EU employs destination-based taxation where the VAT rate of the country in which goods are purchased is applied. Under destination-based taxation, the level of public goods and services in a resident's home jurisdiction is typically lower if the neighboring

[2] This is precisely the situation with regard to the case of Washington with respect to neighboring Oregon and Idaho.

jurisdictions allow for the resident to evade paying higher after tax prices (Bovenberg, 1994). Overall, societal welfare in tax systems where governments set tax rates to fully fund their public expenditure cannot be improved at the macroeconomic level by evasion, and this evasion is Pareto inefficient (Lovely, 1994).

In contrast, origin-based taxation prescribes application of consumers' home country VAT rate to all purchases, regardless of where the sale occurs. Proponents of this system argue that border tax effects would greatly diminish and the adverse trade effects would no longer hamper a country's ability to generate revenue (Lockwood et al., 1994). To advance the political feasibility of such a move, some studies propose that country specific VAT rates also apply to exports to countries outside the EU (Genser, 1996). Furthermore, to maximize the redistribution of tax revenue under origin based taxation, with country specific export rates, Keen and Wildasin (2004) propose an additional source-based tax on capital (intermediate goods). Doing so, according to their model, is Pareto efficient and gains accrue to all countries within the overarching tax system, such as the EU member states.

In the United States, governments in neighboring jurisdictions with different taxing systems face many of the tax revenue challenges that arise internationally. In general, raising taxes is typically considered a political *faux pas* as many residents are opposed to tax increases. However, Luna (2004) finds that local governments usually raise their sales tax in response to a similar measure in a neighboring jurisdiction. When adjusting local sales tax rates, governments are also sensitive to their proximity to large, metropolitan statistical areas (MSA). Counties on the urban fringe of an MSA have a lower ability to raise rates as consumers typically shop in the urban core to take advantage of agglomeration economies and the diversity of retail establishments (Rogers, 2004).[3]

B. Empirical Estimates of Border Effects

The earliest empirical models of the border tax effect examine retail and food sales in jurisdictions along state borders where the presence of a tax differential was believed to impact the distribution of sales. Fisher (1980) estimated the price elasticity of food sales in Washington, D.C. and Fox (1986) examined metropolitan areas along the Tennessee border and the significance of sales and income taxes on retail sales and employment. Following in their footsteps were Walsh and Jones (1988), who analyzed the border tax effect on food sales in West Virginia, as the state phased out its sales tax on food from 1980 to 1982. Using a pooled, multivariate model for all West Virginia counties, Walsh and Jones found that the price elasticity, due to the food sales tax, was -5.9 and statistically significant in the border counties, while no such effect was seen in the interior counties. Tosun and Skidmore (2007) reexamined West Virginia food sales given the state's reintroduction of the food sales tax in 1989 at the high rate of 6%. Included in their estimations was a spatial error correction to incorporate the possibility of a county's sales affecting neighboring counties' sales.

Two previous studies are of particular interest here as they examine the border tax effect in Washington State. Brown (1990) estimated a pooled, cross-sectional model and found the short- and long-run price elasticity to be -1.8 and -2.4, respectively. Beck (1992) incorporated a dummy variable to allow for elasticity differences for counties in or near the Portland-Vancouver

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[3] Levin and Wright (2004) also show that under certain circumstances, it may even be advantageous for retailers to locate their stores in the higher taxing jurisdiction. By adjusting pre-tax prices, such as a car manufacturer may do, under a destination based tax structure, the company effectively bears a portion of the sales tax. Under such a scenario, consumers are more willing to purchase the good in their (higher taxing) home jurisdiction to avoid the costs associated with border crossings (distance, time, gas, etc.).

MSA. Given the 1983 decision of the Washington legislature to increase the sales tax rate in all counties outside the Portland-Vancouver MSA region, this was an appropriate model to test the legislature's reasoning. The author found that the price elasticity for all counties ranges between -2 and -3.2 while the price elasticity for the counties near the Portland-Vancouver MSA was in the -4.1 to -6.1 range.

Finally, event-specific studies have, likewise, used the general market demand for retail goods to estimate the effects of the border tax effect. When two new malls opened in the late 1990s in Tennessee, Chervin, Edmiston, and Murray (2000) found a 15.9% decline in sales in the neighboring counties, regardless of the differing tax rates. This result indicates the drawing power of large shopping destinations to attract customers who prefer shopping diversity and the convenience of one location over shopping outside their home county strictly for the purpose of paying lower, after tax prices. [4]

III. Methodology

A. Model

In much of the literature, estimates of the border tax effect are commonly obtained using a standard Ordinary Least Squares (OLS) or Fixed Effects (FE) regression technique.[5] In order to properly capture the effect of the border on taxable retail sales in Washington, it is imperative to

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[4] Studies on alcohol sales across neighboring jurisdictions also fall in the border tax effect category as alcohol is generally taxed at a wide range of values, which are usually rooted in cultural history. Within the United States, where geographical distances are small (e.g. New England), Beard, Grant, and Saba (1997) find border crossings for alcohol purchases to be a significant factor in the level of sales. Likewise, the opening of the bridge connecting Malmo, Sweden to Copenhagen, Denmark in 2000, coupled with the 2003 Danish VAT rate cut, significantly decreased alcohol sales in southern Sweden (Asplund et al., 2007).

[5] Alterative approaches include, Asplund, Friberg, and Wilander (2007) who use first differencing to control for unobservable factors, and Torralba (2004) who uses instrumental variables to correct for the perceived endogenous nature of the sales tax rate. To check for endogeneity in our analysis, we also conducted an instrumental variable (IV) analysis. Using combinations of population, mileage, percent of youth and/or elderly in the population as instrumental variables we find that the OLS equivalent of our fixed effects estimation is statistically superior to the IV results. Our Hausman test for IV versus OLS with county and time dummies fails to reject the null hypothesis that OLS is consistent and IV is inappropriate. Results are available upon request.

correctly specify the market demand. In doing so, by controlling for key economic factors affecting consumer spending, the model can quantifiably describe the magnitude the border has on retail sales, which directly impacts state and local governmental jurisdictions' budgets and fiscal decisions.

We specify general market demand for taxable retail goods in two ways. First, following Brown (1990), the demand is estimated in a semi-log format, as shown in Equation (1). Second, following Beck (1992), the demand is estimated in a log-log format, providing desired elasticity measures as in Equation (2) below:

$$\begin{aligned} sales_{it} &= \beta_0 + \beta_1 \ln price_{it} + \beta_2 \ln(price_{it} * border_i) + \beta_3 \ln income_{it} \\ &\quad + \beta_4 \ln travel_{it} + \beta_5 X_{control} + \varepsilon_{it} \end{aligned} \tag{1}$$

$$\begin{aligned} \ln sales_{it} &= \beta_0 + \beta_1 \ln price_{it} + \beta_2 \ln(price_{it} * border_i) + \beta_3 \ln income_{it} \\ &\quad + \beta_4 \ln travel_{it} + \beta_5 X_{control} + \varepsilon_{it} \end{aligned} \tag{2}$$

where *i* denotes counties, *t* denotes years, and ε is a zero mean, white noise error term. In both models the independent variables remain the same. The dependent variable in Equation (1) is real per capita taxable retail sales measured in levels. This model allows us to obtain estimates to generate revenue gains to Washington State from complete or partial tax harmonization. Using the natural log of real per capita taxable sales as a dependent variable in Equation (2) allows us to obtain the desired price elasticity measure.

Consistent with previous literature (see for example, Tosun and Skidmore, 2007), explanatory variables in Equations (1) and (2) include: real per capita income (*income*), a county relative price level measure and a travel cost proxy. Income is expected to be positive and statistically significant. A standard measure of the county relative price, *price*, is the ratio of after-tax prices in a resident's home county (H) relative to a neighboring county (N):

$$price = \frac{P_H(1+t_H)}{P_N(1+t_N)} \quad (3)$$

Equality between pre-tax prices in home and neighboring counties can be assumed if factor prices are the same in these jurisdictions and goods are sold in competitive markets. If $P_H = P_N$ in Equation (3), a standard assumption we adopt here, then the price differences that consumers face will fully reflect the tax difference. If not we believe that price will be higher in Washington but by less than the full amount of the tax (Levin and Wright, 2004). The relative price measure is expected to be largest for counties along the Oregon-Washington border where the denominator is equal to 1 due to the lack of a sales tax in Oregon. Overall, we expect *price* to have a significant and negative effect on retail sales, indicating the higher the relative after tax home price, the more likely consumers are to evade paying higher prices due to the sales tax. To capture the border effect in Equations (1) and (2), we include an interaction term of the price variable with *border*, a dummy variable that takes on the value of 1 if the observed county is on the border with either Oregon or Idaho, and zero otherwise. The sum of the estimated coefficients on *price* and the corresponding interaction term in Equation (2) generates the desired price elasticity measure which is expected, according to estimates in previous literature, to be between -2 and -11.

Finally, is it standard to include a variable that serves as a proxy for travel costs. Our measure, *travel*, is intended to capture travel costs that consumers will, on average, incur if they choose to shop in jurisdictions other than their home county. As a proxy for travel costs we use the distance (in miles) between county seats[6] adjusted by the percent change in the gasoline price

[6] Border counties' seats were measured to the closest Oregon or Idaho county, while all other Washington counties were measured from their county seat to either Portland, OR or Coeur D'Alene, ID, whichever is closest. Portland and Coeur D'Alene were chosen as the most logical, large retail locations close to the border in which consumers may choose to purchase goods. In doing so, the expected sign of travel costs remains positive (+) and the model is corrected to avoid a myriad of 0 data points. For four counties, county seat distances were inappropriate to model

index (GPI) created by the United States Energy Information Administration over the sample period.[7] As shown in Equation (4), distance (*Dist*) is multiplied by the ratio of GPI in year t to GPI in 2006 (t = 1992, …, 2006). This adjustment allows the *travel* variable to capture relative fuel price fluctuations over sample years and their impact on consumer behavior.

$$Travel = Dist\left(\frac{GPI_t}{GPI_{2006}}\right) \tag{4}$$

Travel is expected to have a significantly positive effect on taxable retail sales, indicating that the further consumers are required to travel to shop in alternative jurisdictions, the more likely they are to purchase goods in their home county. Additional control variables in both models are represented as an nxm matrix, $X_{Control}$. In previous literature (see for example Brown, 1990; Beck, 1992), it is common to control for key demographic, economic and geographic determinants pertaining to retail sales. We include as controls the county level unemployment rate, the percentage of the population that is 65 years and older, the percentage of the population that is 18 years and younger, the number of retail establishments per 1,000 residents, as well as year dummies. The unemployment rate is expected to have a negative effect on the level of sales as, theoretically, the larger the number of unemployed individuals, the lower the income and consequently, expenditures on consumer goods.

The elderly and youth percentages are used to control for the demographic characteristics at the county level. Theoretically, the elderly spend a sizable share of their income on non-taxable goods such as food, housing and medical care, which would decrease the level of taxable retail sales in the county. The percentage of youth is expected to be positively correlated with

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consumer behavior and the distances used by Brown (1990) were used instead. These include Benton County (Tri-Cities, WA to Herminston, OR), Pend Oreille County (Newport, WA to Oldtown, ID), Walla Walla County (Walla Walla, WA to Milton-Freewater, OR) and Whitman County (Pullman, WA to Moscow, ID).

[7] Specifically, the West Coast gasoline price index was used in this study.

sales as parents spend a larger share of their income on retail goods to provide for their children. The number of retail establishments in a county is used to measure the general availability of retail products in the county. Computing the number of stores per 1,000 residents generates the desired measure of retail density which is expected to positively affect retail sales. That is, the more stores available, the more likely residents are to shop in their home jurisdiction. Overall, previous literature has found mixed results pertaining to the above control variables and their statistical significance in the models. Table 1 presents a summary of the dependent and independent variables considered in Equations (1) and (2) along with their sources and expected signs.

Equations (1) and (2) are estimated using fixed effects to control for unobservable county-specific characteristics. Our choice of fixed over random effects estimation is both intuitive and statistically valid. That is, we treat county-specific heterogeneity as fixed since the set of counties included in our sample is predetermined (i.e. we only focus on the 39 counties in Washington State). The alternative, that county-specific effects are random, would apply if counties were drawn at random from a larger sample (for example all counties in the U.S.). Our Hausman test for fixed versus random effects overwhelmingly rejects the null hypothesis that random effects estimation is appropriate.[8]

To further evaluate the robustness of the estimates produced by equations (1) and (2) we also conduct a spatial autocorrelation analysis. In recent literature (see for example Blonigen et al., 2007; Tosum and Skidmore, 2007), the increased emphasis on empirically modeling spatial interactions is justified by the consequences that omitting spatial autocorrelation corrections can have on the validity of the estimated results. Specifically, in the presence of spatial

[8] The Hausman test for fixed versus random effects produces a Chi-squared statistic of 102.003 (Probability = <0.0000). We reject the null hypothesis that that the coefficients estimated by the efficient random effects estimator are the same as the ones estimated by the consistent fixed effects estimator. Results are available upon request.

autocorrelation (as can be suspected in panel data sets where observations are characterized by geographical proximity), OLS or FE estimation techniques may produce estimates that are biased, inefficient and/or inconsistent. Since our analysis concerns counties that are spatially related, it is necessary to evaluate our fixed effects estimation results relative to those obtained from a specification that controls for spatial dependence.[9] To this end, we estimate equations (1) and (2) as a spatial autoregressive model (SAR) using a maximum likelihood approach (Blonigen et al., 2007). In particular, we estimate:

$$Sales\ Measure_{it} = \beta_0 + \beta_1(Demand\ Variables) + \rho \cdot W \cdot Sales\ Measure_{it} + \varepsilon_{it} \tag{5}$$

For comparison purposes, Equation (5) is estimated using real, taxable retail sales per capita as the dependent variable, and separately using the natural log of this measure. Along with the covariates and controls from Equations (1) and (2), *Demand Variables*, Equation (5) includes a spatially lagged dependent variable, *W* · *Sales Measure*, where *W* is a row-standardized, symmetric, contiguity matrix that parameterizes the distance between neighboring counties[10]. The weighting matrix *W* and its elements are defined as follows:

$$W = [w_{ij}] \quad and \quad w_{ij} = \begin{cases} d_{ij} \ if \ i \ and \ j \ are \ neighbors; \\ 0 \ otherwise \end{cases} \tag{6}$$

where d_{ij} is a weight as each row of *W* is normalized so it sums up to unity. The coefficient ρ measures the dependence of sales in county *i* on proximity to its neighbors. If spatial dependence exists in our data, the estimated coefficient on ρ is expected to be statistically significant, indicating that proximity of the observed county to its neighbor(s) significantly explains the

[9] We also estimated a spatial error model (SER), which corrects for correlation between neighboring counties' error terms. The results are qualitatively the same as those from the SAR model reported in Table 3. This is consistent with results from diagnostic tests in Tosun and Skidmore (2007) who find no evidence of spatial dependence in either the SAR or SER specifications for the case of West Virginia.

[10] Construction of the weighting matrix follows the convention of using queen contiguity based on geographical county borders. In unreported results, we also use an alternative definition of the weighting matrix based on major highways/routes as a proxy for county contiguity. Empirical results using this alternative remain qualitatively the same and are available upon request.

variation of sales in that county. In other words, the spatial autocorrelation term in Equation (5), allows for sales in county *i* to depend on a weighted average of its neighbors' sales.

B. Data

The present study focuses on Washington State's 39 counties over the years 1992 to 2006 (inclusive). For all variables detailed in Table 1, data is available for the entire 15 year period resulting in a panel of 585 observations. We obtain data on county tax rates from the Washington State Department of Revenue.[11] This is also the source for information on real, per capita taxable sales by county. Information on demographic variables such as persons 18 years or younger and 65 years or older is available from the Washington State Office of Financial Management. Our measures of the county-specific unemployment rate and per capita personal income are collected from the U.S. Bureau of Economic Analysis. Finally, information on retail establishments per 1000 people was obtained from the Washington State Employment Security Department. As shown in the descriptive statistics in Table 2, there is a large and statistically significant difference in all variables between border and interior counties in Washington. Difference-in-means tests for all key variables, such as per capita retail sales and the *price* measure, are significant at the 1% level. Thus the univariate analysis is suggestive of a border effect, an issue we investigate further in our fixed effects estimation below.

IV. Empirical Results

A. Fixed Effects Estimates and Spatial Dependence Analysis

Table 3 presents the empirical results from estimation of Equations (1) and (2). For comparison purposes, we provide estimates obtained using fixed effects and maximum

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[11] Washington state's portion of the sales tax has remained constant at 6.5% over the sample period with local jurisdictions' rates typically increasing over time. Idaho state's sales tax rate has increased from 5% to 6% in 2003, then decreased back to 5% in 2005 before increasing to 6% in 2006. Oregon state's sales tax as remained 0% over the sample period. In total, Washington's sales tax rates have increased over the sample period while the neighboring states' have remained, more or less, constant.

likelihood estimation with a spatial lag variable (SAR) side-by-side. Our semi-log specification is tabulated as Model (1) and our double-log specification as Model (2). Focusing on the fixed effects results first, we find that overall, the estimated coefficients on the traditional determinants of demand for taxable retail goods are reasonable, significant, and have the expected signs. In both models, higher income per capita and higher travel costs are, on average, significantly related to higher per capita taxable sales in the observed county. A one percentage point increase in per capita income translates into $100.48 increase in real per capita sales while Model (2) indicates a retail sales income elasticity of 0.856. The latter suggests that for each additional dollar of income, an individual spends approximately 86 cents on taxable retail goods, a result that is generally in line with macroeconomic data at the national level.[12]

The estimated coefficients on the county relative price measure and its corresponding interaction term with the *border* dummy are significant throughout, and together, provide an overall estimate of the effect of county relative price on taxable retail sales. After adjusting for units of measurement, we estimate for Model (1) that the overall effect of a one percentage point increase in the county relative price is a $333.11 decrease in real per capita sales at the county level. More importantly, the negative and significant coefficient on the price measure interacted with the border dummy illustrates the border sales tax effect due to the after tax price discrepancy between counties in Washington and neighboring, lower tax jurisdictions in Idaho and Oregon. That is, given the large discrepancy in sales tax rates between Idaho, Oregon and Washington, the availability of lower priced goods is easily obtainable by the nearly 20% of Washingtonians who live in border counties. Using the estimates in Model (2), we find the elasticity of the price measure to be -3.11. Brown (1990) found the long run price elasticity of

[12]On the national level, using data from the United States Bureau of Economic Analysis, personal consumption expenditures account for between 79.2% and 84.8% of total personal income over the years 1992-2006.

Washington border counties to be -2.4 and Beck (1992) found it to be between -2 and -3.2 given the different models estimated. The results of this study are thus in line with those obtained in previous literature.

The travel cost estimator in Models (1) and (2) has the expected positive sign, indicating that after adjusting for the cost of gasoline, the further a resident must travel to a lower taxing jurisdiction, the higher the level of expenditures on retail goods are in the home county. Interestingly, unlike previous literature where this variable does not significantly explain variation in retail sales (see Walsh and Jones, 1988; Brown, 1990), we find our measure of travel costs to be highly significant in both models. We attribute this significance to the way we construct the variable, namely adjusting distance by the gasoline price index and then using the most relevant index in this category, pertaining to West Coast prices.

Several observations regarding the remaining socio-economic control variables are of interest.[13] The percentage of a county's population that is 18 years or younger is highly significant and positively related to taxable retail sales. This is consistent with our expectation that parents will, on average, spend a larger portion of their income on retail goods to provide for their children. Also as expected, the percentage of elderly in the population has a negative and significant effect on retail sales. Individuals over 65 are usually retired from the labor force and spend the majority of their income on non-taxable goods, such as food, housing and medical care Finally, the number of retail establishments per 1,000 people in a county, used as a proxy for shopping availability, has a strong, positive effect on taxable retail sales level indicating that greater availability of retail outlets in a resident's home county is associated with higher sales in that county.

[13] We note that county-level unemployment rate is insignificant throughout and that its behavior in our analysis is similar to that documented by Tosun and Skidmore (2007).

Turning to our results from analysis of spatial dependence in our sample we note two important observations related to the maximum likelihood estimation of Models (1) and (2) using a spatial lag model. First, the estimated coefficients on key determinants of the demand for taxable retail goods remain remarkably stable which implies that county-specific heterogeneity is fixed and properly accounted for by the fixed effects technique. This is both intuitive and in line with previous literature (see for example Blonigen et al., 2007; Garretsen and Peeters, 2007). Second, the point estimates of the spatial terms from the SAR estimation of both Models (1) and (2) are highly insignificant and very small in magnitude. This is also intuitive since spatial characteristics among counties in our sample are fixed over time. Thus, the use of county-level fixed effects (or county dummies in an OLS equivalent of the fixed effects estimation) eliminates the statistical and economic significance of the spatial terms (*W · Sales Measure*). Overall, we conclude that there is no significant spatial dependence in our sample and that the fixed effects results presented above are robust to both heteroscedasticity and (spatial) autocorrelation corrections.

B. Estimates of Revenue Gain from Full and Partial Tax Harmonization

A key motive behind the empirical analysis in this paper is to provide an estimate of the gain in sales tax revenue from complete or partial elimination of differences in border county relative prices due to sales taxes. This topic is of primary importance for fiscal policy at the state and local level due to the financial impact the border tax effect has in Washington State. Using total taxable retail sales in 2006 and our overall estimate of the county relative price in Model (1), we calculate the gains from elimination of relative price discrepancies in border counties and present these by county. [14] Table 4 presents these estimates by county. Overall, we find that the

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[14] For border counties in Washington State, elimination of county relative price differences due to lower tax rates in neighboring Oregon and Idaho can result from Washington lowering its sales tax rate, Idaho and Oregon raising

estimated gain in taxable retail sales from the elimination of relative price differences in each of the fourteen border counties amounts to approximately $2.2 billion in 2006, or nearly 13% of sales in 2006. This corresponds to over $145 million in tax revenue at the state level where the tax rate is 6.5%. Additionally, at the local (county) level, complete sales tax harmonization would result in over $21 million dollars, which could be used to fund public goods and services.

While the magnitude of the tax revenue generated through sales tax harmonization between Washington State and its neighbors is astounding, as seen in the review of theoretical considerations in Section II, such a policy is Pareto inefficient. Similar to a binding minimum tax, Table 5 calculates the estimated gains in taxable retail sales and subsequent tax revenue from a 1% decrease in the relative county price in border counties due to a sales tax discrepancy. As before, a one percentage point reduction in the county relative price would most realistically be obtained as a result of tax increases in Oregon and Idaho. Enacting such policy changes would produce an estimated increase in retail sales in Washington's border counties of over $434 million, based on 2006 figures. These additional sales would generate over $28 million in state level tax revenue and nearly $5 million at the local or county level.

V. Conclusion

The purpose of this paper has been to quantifiably determine the effects of the border tax problem between neighboring jurisdictions and particularly focusing on the case of Washington State. We evaluate the border tax effect using data on the determinants of taxable retail sales from 1992 – 2006 using a fixed effects estimation. Our results suggest that the price elasticity arising from the discrepancy in sales tax rates between Washington, Idaho, and Oregon is -3.11 and this estimate falls in the range expected from estimates in previous literature. The cost of the

theirs, or, with highest probability given the current political and economic climate, a combination of the these alternatives where Washington and Oregon (and/or Idaho) meet in the middle.

sales tax discrepancy to Washington State is estimated to be nearly $2.2 billion in sales or over $145 million in tax revenue at the state level and approximately $21 million at the local level. Furthermore, our analysis of spatial dependence in our sample reveals that spatial autocorrelation is not an issue for our estimated results which remain remarkably stable in maximum likelihood estimations that include a spatial lag term.

While the present study builds and furthers previous literature in determining the border effect arising from neighboring jurisdictions with different tax structures, there remains a need to incorporate the changing dynamics of consumer behavior. Possibly the most important aspect with regards to retail sales in the past decade has been the rapid development and use of the internet. Continuing the 1998 Internet Tax Freedom Act, President George W. Bush signed the Internet Tax Act Freedom Amendments Act in 2007, which extends the moratorium on electronic commerce taxation. Given the physical presence rule for taxing businesses, the internet provides an increasingly accessible outlet for consumers to evade the sales tax. This directly affects governments' ability to effectively plan and pay for the requisite public goods and services for their citizens.

In future research it will be of interest to reexamine the border tax effect in light of recent sales tax policy initiatives in Washington State. Specifically, in July of 2008, Washington implemented a sales tax policy change that is a move toward origin-based taxation in the nomenclature of previous literature. Under the new provisions, the sales tax on goods which are home delivered is charged based on the jurisdiction in which delivery occurs, not that of the retailer. The motive behind this legislative change is to reduce the border tax effect and redistribute tax revenue within the state. While this marks a step in the theoretically supported direction, it only applies to Washington-based retailers and therefore, cannot reach across state lines to reduce the true border tax effect. While at present only two quarters worth of data exist under the new tax regime, the effectiveness of this policy in minimizing the border tax effect would warrant reexamination in future analyses that can benefit from a longer time span.

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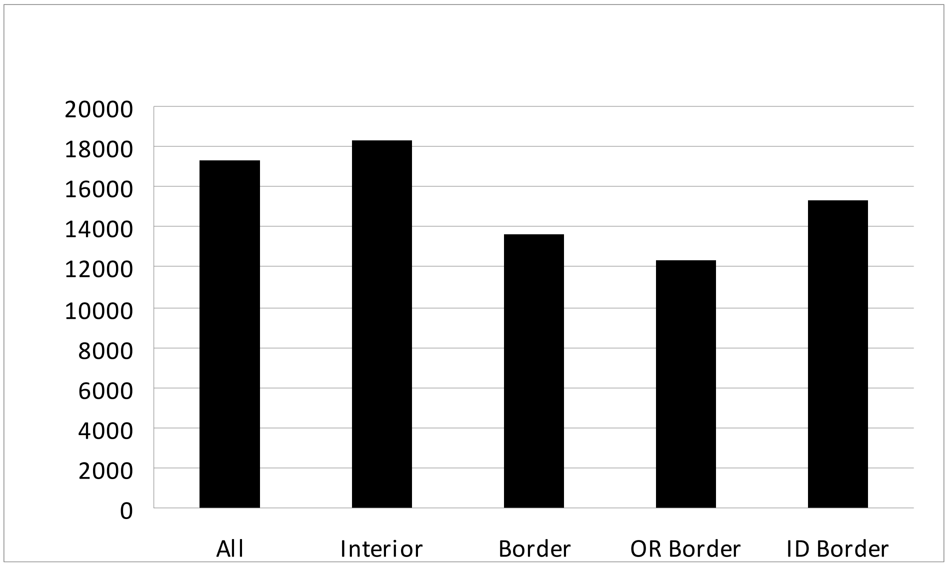
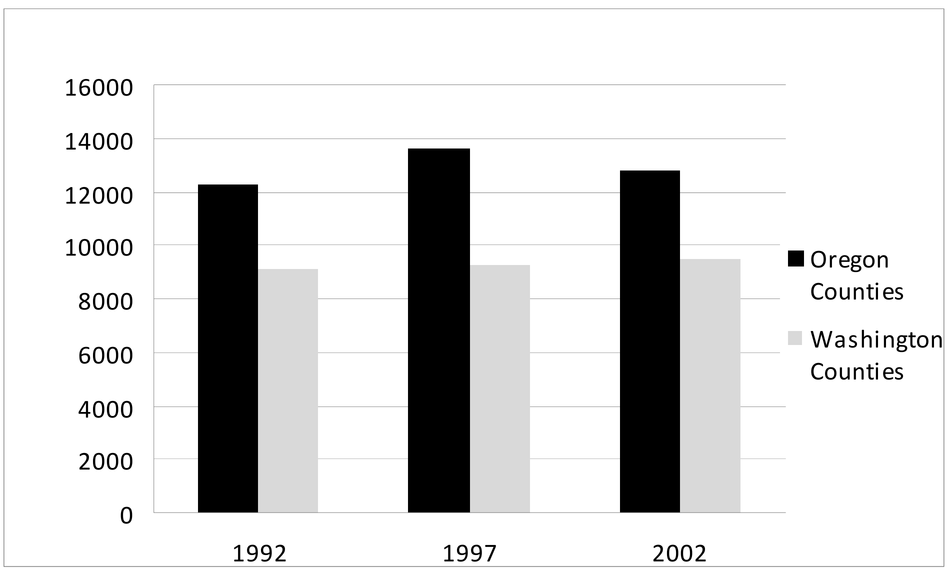
Figure 1. Per Capita Sales in Washington Counties in 2006



Source: Washington State Department of Revenue, 2006

Figure 2. Washington-Oregon Border Counties: Real Per Capita Retail Sales



Source: U.S. Census Bureau's Economic Census (1992, 1997, 2002). Oregon Counties: Clatsop, Columbia, Gilliam, Hood River, Morrow, Multnomah, Sherman, Umatilla, Wallowa, Wasco. Washington Counties: Asotin, Benton, Clark, Columbia, Cowlitz, Garfield, Klickitat, Pacific, Skamania, Wahkiakum, Walla Walla

Table 1. Variable Definitions, Sources, and Expected Signs

Variable	Description	Source	Expected Sign
Sales	Per capita real taxable sales in dollars.	Washington State Department of Revenue	NA
Income	Per capita real personal income in dollars.	U.S. Bureau of Economic Analysis	+
Price	County relative price reflecting differences in home county tax rate relative to neighboring county tax rate	Computed by the authors; County tax rate source: Washington State Department of Revenue	-
Travel	Mileage distance between county seats of neighboring jurisdictions, adjusted by the West Coast Gasoline Price Index	Computed by the authors; Gasoline Price Index source: United States Energy Information Administration	+
Unemployment	County-specific unemployment rate in percent.	U.S. Bureau of Labor Statistics	-
Youth	Persons 18 years and younger, (percentage of county population)	Washington State Office of Financial Management	+
Elderly	Persons 65 and older (percentage of county population)	Washington State Office of Financial Management	-
Retail Establishments	Number of Retail establishments per 1000 residents	Washington State Employment Security Department	+

Table 2: Descriptive Statistics

Variable	Overall		Interior Counties		Border Counties	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Sales (dollars/capita)	11,612	4,161	12,843	4,179	9,413***	3,090
Inc (dollars/ capita)	28,060	5,491	28,840	6,343	26,666***	3,035
Price (index)	1.020	0.031	1.002	0.002	1.059***	0.024
Travel (index)	67.424	52.519	95.245	45.231	17.742***	13.310
Unemp (percent)	7.377	2.479	7.425	2.215	7.292	2.895
Youth (percent)	27.691	3.726	28.063	4.103	27.027***	2.823
Elderly (percent)	14.078	3.700	13.680	3.432	14.790***	4.047
RetailEst (per 1000 people)	3.680	1.132	3.880	1.188	3.315***	0.921
N	585		375		210	

***, **, and * indicate significance of difference-in-means test of interior versus border counties at the 1%, 5%, and 10% levels, respectively.

Table 3: Border Effects – Fixed Effects (FE) Estimation and Spatial Lag (SAR) Analysis

	Model (1): Semi-Log Specification; Dependent Variable = $Sales_{it}$		Model (2): Double-Log Specification; Dependent Variable = $\ln(Sales)_{it}$	
	FE	**SAR**	**FE**	**SAR**
Intercept	-117268.773*** (18,622.741)	-110203.902*** (17,438.580)	-3.094* (1.591)	-2.474 (3.460)
Ln(Real, Per Capita Income)	10048.155*** (1003.672)	10048.155*** (951.621)	0.856*** (0.090)	0.856*** (0.085)
Ln(County Relative Price)	71078.396** (32193.261)	71078.406** (30523.773)	6.730*** (2.182)	6.730*** (2.069)
Ln(CountyPrice*Border)	-104389.464** (40371.182)	-104389.478*** (38277.613)	-9.843*** (3.397)	-9.843*** (3.221)
Ln(Travel)	2455.594*** (370.323)	181.044 (163.513)	0.249*** (0.035)	0.036* (0.019)
Ln(Unemployment Rate)	-83.902 (208.637)	-83.902 (197.818)	0.028 (0.021)	0.028 (0.020)
Ln(Youth Percentage)	8678.526*** (2748.347)	8678.528*** (2605.823)	1.160*** (0.265)	1.160*** (0.251)
Ln(Elderly Percentage)	-4956.940* (2775.291)	-4956.939* (2631.368)	-0.499* (0.262)	-0.499** (0.249)
Ln(Retail Establishments)	1614.502*** (471.628)	1614.502*** (447.170)	0.143*** (0.048)	0.143*** (0.046)
Spatially Weighted Retail Sales ($W \cdot Sales\ Measure_{it}$)		3.80e-08 (0.339)		5.54e-08 (0.339)
Time Dummies	Yes	Yes	Yes	Yes
Number of Observations	585	585	585	585
R^2 / Log-Likelihood	0.3547	-4742.84	0.3193	694.08

Robust standard errors are in parentheses. ***, **, and * indicate coefficient estimates are significantly different from zero at the 1%, 5%, and 10% levels.

Table 4: Estimated Gain in 2006 Taxable Retail Sales from Elimination of Tax Differential

County	Total Taxable Retail Sales (2006)	Estimated Gain in Taxable Retail Sales	Estimated State Tax Revenue	Estimated Local Tax Revenue
Asotin	$183,624,442	$6,614,761	$429,959	$33,074
Benton	$2,303,245,278	$404,105,660	$26,266,868	$4,849,268
Clark	$4,866,777,344	$967,864,073	$62,911,165	$5,807,184
Columbia	$29,770,738	$10,431,561	$678,051	$146,042
Cowlitz	$1,337,394,181	$246,912,462	$16,049,310	$2,469,125
Garfield	$15,899,676	$5,291,463	$343,945	$52,915
Klickitat	$162,750,735	$46,204,117	$3,003,268	$231,021
Pacific	$195,060,498	$53,466,762	$3,475,340	$534,668
Pend Oreille	$89,831,028	$6,367,568	$413,892	$70,043
Skamania	$87,112,482	$24,539,598	$1,595,074	$122,698
Spokane	$7,278,765,098	$281,190,667	$18,277,393	$4,217,860
Wahkiakum	$24,290,624	$9,865,906	$641,284	$98,659
Walla Walla	$718,942,577	$153,521,385	$9,978,890	$2,302,821
Whitman	$410,491,705	$23,420,572	$1,522,337	$304,467
Total	$17,703,956,406	$2,239,796,555	$145,586,776	$21,239,844

Table 5: Estimated Gain in 2006 Taxable Retail Sales from 1% Decrease in Tax Differential

County	Total Taxable Retail Sales	Estimated Gain in Taxable Retail Sales	Estimated State Tax Revenue	Estimated Local Tax Revenue
Asotin	$183,624,442	$6,217,875	$404,162	$31,089
Benton	$2,303,245,278	$52,481,255	$3,411,282	$629,775
Clark	$4,866,777,344	$136,318,884	$8,860,727	$817,913
Columbia	$29,770,738	$1,320,451	$85,829	$18,486
Cowlitz	$1,337,394,181	$32,921,662	$2,139,908	$329,217
Garfield	$15,899,676	$705,528	$45,859	$7,055
Klickitat	$162,750,735	$6,600,588	$429,038	$33,003
Pacific	$195,060,498	$7,128,902	$463,379	$71,289
Pend Oreille	$89,831,028	$4,218,514	$274,203	$46,404
Skamania	$87,112,482	$3,505,657	$227,868	$17,528
Spokane	$7,278,765,098	$149,031,053	$9,687,018	$2,235,466
Wahkiakum	$24,290,624	$1,315,454	$85,505	$13,155
Walla Walla	$718,942,577	$19,190,173	$1,247,361	$287,853
Whitman	$410,491,705	$13,792,115	$896,487	$179,297
Total	$17,703,956,406	$434,748,110	$28,258,627	$4,717,530